Do MSRPs Decrease Prices?*

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Abstract

The nature of manufacturer's suggested retail prices (MSRPs) and whether their effect is pro- or anticompetitive is not well understood. We exploit a policy experiment in which a ban on MSRPs was imposed and then lifted a year later. Prices increased by 2.3 percent as a result of the ban and decreased by 2.6 percent when the ban was lifted. We find no indication that MSRPs lowered prices by acting as binding price ceilings and outline an alternative mechanism in which recommendations affect prices indirectly by providing information to searching consumers. We demonstrate that recommendations can increase search and reduce prices.

Keywords: recommended retail price, suggested retail price, list price, non-binding price, search with uncertainty, vertical restraints, resale price maintenance

JEL Classification Numbers: L110, L400, L810

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1 Introduction

Manufacturers often attempt to exert influence over the prices set by retailers. Whether such practices are pro- or anticompetitive is the subject of an active academic and policy debate.[1] Evidence of this debate is reflected in the disparate treatment of the presumption of illegality of vertical price restraints in antitrust regulation over time and across countries and jurisdictions.[2] The debate persists in part due to the existence of competing theoretical models that lend credence to both sides of the argument and a scarcity of empirical evidence to determine which theories are supported.[3]

In this paper we estimate the effect of a common vertical price restraint, the manufacturer's suggested retail price (MSRP), by exploiting a natural experiment in which a ban on MSRPs was initially imposed and then lifted a year later. Although MSRPs are used in a variety of consumer products, ranging from new cars, electronics, and appliances to inexpensive items such as books and packaged foods, the mechanism by which they affect prices is not well understood. Specifically, it is unclear whether manufacturers make recommendations in order to increase or reduce prices.

The literature on vertical restraints provides several explanations for why a manufacturer may want either higher or lower retail prices. Higher prices and thus higher downstream profits can better align the incentives of retailers to those of the manufacturer, for instance with respect to the provision of quality-enhancing services (Telser, 1960; Klein and Murphy, 1988). Higher retail prices may also be part of a collusive upstream agreement, in which vertical price restraints keep manufacturers from engaging in secret wholesale price cuts (Jullien and Rey, 2008). In addition, higher prices may be the objective of a manufacturer that acts as a cartel leader for powerful retailers. A manufacturer may likewise implement policies to reduce prices. For example, when

[1] See O'Brien (2008) for a review of competing theories of vertical restraints since Cournot's (1838) seminal article.

[2] In the U.S. the treatment of vertical restraints has changed in the past three decades, relaxing the previous *per se* prohibition. At the federal level, the 1985 Vertical Restraints Guidelines issued by the Department of Justice were permanently withdrawn in 1993. Following the U.S. Supreme Court's decision in *Leegin*, issued in 2007, all vertical restraints are evaluated under the rule of reason approach.

[3] Lafontaine and Slade (2008) summarizes the current body of empirical evidence regarding the effects of vertical restraints.

retailers have market power they impose an additional margin above the cost they pay to the manufacturer. This results in lower sales than the manufacturer prefers, an effect known as double marginalization (Cournot, 1838; Spengler, 1950; Mathewson and Winter, 1984).

These different manufacturer motives have varying implications for welfare. If MSRPs are used to overcome externalities in quality provision or to reduce double marginalization, they likely improve welfare. However, if MSRPs are used to facilitate upstream or downstream collusion, their effect is anticompetitive. Empirical work is then necessary to discern which of the competing theories of vertical price restraints is most relevant in explaining manufacturer price recommendations.

A challenge in estimating the effect of MSRPs, and of vertical price restraints in general, is that their use is endogenous to industry characteristics which are often unobserved (Lafontaine and Slade, 2008). In particular, the use of MSRPs can be endogenous to unobserved conditions that determine the vertical relationship itself. We are able to address the issue of endogeneity by exploiting an unusual natural policy experiment in South Korea. In Section 2 we outline the policy imposed by the Korean government in July 2010, which banned MSRPs on products in several processed foods categories with the stated purpose of fostering competition and reducing prices. One year later, in response to public pressure from consumers, the government lifted the ban. The ban and its subsequent lift afford us the unique opportunity to first estimate the effect of MSRPs when they are removed and then validate these results by observing this experiment in reverse.

Our data comes from the Nielsen Korea Retail Measurement Service and is described in Section 3. This data set contains monthly prices and sales quantities for 253 products, including products subject to the ban and other similarly priced products that remained outside the scope of the regulation and serve as a suitable control group. The empirical results, presented in Section 4, employ a difference-in-differences approach to separately estimate the price effect of the MSRP ban and of its reversal. Contrasting the two effects allows us to assuage concerns about unobserved factors affecting the treatment group contemporaneously with the ban or with its reversal. We find that prices rose by 2.2 percent as a result of the ban and fell by 2.6 percent as a result of the lift of the ban. The two effects are of similar magnitudes and remain so for different specifications of treatment and control groups.

The finding that MSRPs reduced prices suggests that fostering collusion or inducing higher quality provision were not the manufacturer's primary motives. On the other hand, our finding is consistent with the manufacturer attempting to reduce double marginalization, and in Section 5 we explore this idea further. One potential mechanism by which MSRPs reduce double marginalization is by acting as price ceilings. This could occur either if the manufacturer punished retailers for non-adherence or if consumers refused to pay above the recommended price. A closer inspection of the distributions of retail prices casts doubt on this explanation. Most retailers charged prices not just below, but significantly below the MSRP before and after the ban and after the lift of the ban. Furthermore, there is little evidence of retailers raising prices from below to above the MSRP level after the ban, or reducing prices from above to below the MSRP level after the ban's reversal. Although price recommendations impacted prices overall, they did not act as binding constraints.

We propose an alternative explanation of MSRPs as information which helps consumers evaluate retailers' offers. A consumer who decides whether to accept the price at the current retailer or to engage in costly search uses the manufacturer's recommendation to form an expectation of the prices charged by other retailers. Consumers benefit from the information conveyed by an MSRP by making more informed search decisions. As has been shown in the literature, manufacturers can also benefit when consumers are thus informed.[4] In contrast to a price ceiling which binds only at high prices, by providing information relevant for search an MSRP can shift demand and have an effect at every price.

When MSRPs provide information, their removal increases consumers' price uncertainty and we analyze how this uncertainty affects consumer search. We describe two potentially countervailing effects of removing price recommendations, building on work by Rothschild (1974) and Bénabou

[4] Janssen and Shelegia (2013) argue that in a search environment the manufacturer is better off when his wholesale price is observed by consumers. Lubensky (2011) demonstrates that even if the manufacturer cannot exogenously commit to truthfully revealing his wholesale price, he may still be able to do so via cheap talk.

and Gertner (1993). First, when recommendations are removed a consumer faces more variance in prices, which increases the option value of searching. Second, in absence of MSRPs, a consumer who observes a high price may infer that prices are high everywhere and thus may be dissuaded from searching. When the latter effect dominates, a ban of MSRPs results in less search, softening competition and consequently increasing prices.

In Section 6, we provide a discussion of the main results. Our theory predicts that MSRPs, even when non-binding, affect prices. Empirically we find a small but significant effect of approximately two percent. More generally, by demonstrating that prices are lower in the presence of MSRPs, we provide evidence that manufacturer attempts to influence the downstream market can be pro-competitive. The Korean ban was intended to reduce prices but had the opposite effect, contributing to its reversal a year later. This highlights the inherent complexity of vertical relationships and the importance of understanding the mechanism behind a vertical restraint in order to construct an effective regulatory policy. Our empirical analysis demonstrates the existence of pro-competitive motives for vertical restraints, yet does not rule out the possibility of anti-competitive motives in other settings. Even with a model of MSRPs as information, our theory predicts that prices can increase in the presence of recommendations given different market conditions.

2 Background

The regulation of MSRPs in Korea has a long history and encompasses a wide range of products. In 1979, the Ministry of Commerce, Industry, and Energy (MCIE) mandated the implementation of factory price labeling for 45 consumer electronics and clothing categories. For other products, use of MSRPs was at the discretion of the manufacturer.[5] MSRPs were included on products such as consumer electronics, clothes, and processed foods. However, by the late 1990's, the Korean government's view of MSRPs had changed. In 1997, the Ministry of Health and Welfare banned

[5] One exception is prescription drugs, for which Article 58 of the Pharmaceutical Affairs Law required pharmaceutical companies to provide an MSRP.

MSRPs on cosmetics, hygiene items, and prescription drugs. In 1999, MSRPs were banned from consumer electronics and durable goods—TVs, VCRs, corded telephones, stereos, and washing machines—and several clothing categories, including suits, children's wear, and sportswear.

Open Price Regulation

In July 2009, with the objective of fostering competition, the MCIE announced the open price regulation, banning MSRPs from several processed foods categories including biscuits and pies, ice cream, ramen, and snacks. In order to give manufacturers time to comply, the ban was set to start one year after the announcement, in July 2010.

A surprising aspect of this regulation is that it was reversed after only one year. In June of 2011, the MCIE announced that it would allow the reinstatement of MSRPs as early as August 2011. Two factors might have influenced the government's decision to lift the ban for these product categories. First, there was a general perception that prices for these food items increased after the ban.[6] Second, a March 2011 survey by the Korean Consumer Agency—a public institution that defends consumer interests in Korea—found that retailers did not consistently display prices of several products that were subject to the open price regulation (e.g. 21.5 percent of retailers did not display prices for ice cream, 48.8 percent for ramen, and 61.2 percent for snacks). In addition, 93.4 percent of consumers surveyed responded that they felt uncomfortable without MSRP information.

Two key aspects of the lift of the MSRP ban should be noted. First, in part due to the short time between the announcement and the lift of the MSRP ban, producers did not immediately include MSRPs on all products by August 2011. For instance, Korea's *Financial News* reported on August 2, 2011 that producers of snacks, biscuits and pies, and ice cream were expected to include recommendations in September when existing product inventories were depleted. A similar article

[6] ChosunBiz.com reported on June 15, 2011 that prices of snacks, ramen, and ice cream increased sharply after the open price regulation. They reported the price of saewookkang (a snack) increased 16 percent at hypermarkets, 12.5 percent at convenience stores, and 10.4 percent at large chain supermarkets. Similarly, HyundaiCapital.com reported on July 7, 2011 that negative publicity over the open price regulation was mainly due to soaring prices and confusion from the absence of MSRPs. They reported that prices had grown in the past year well above the CPI increase of 3.8 percent, in particular 7.8 percent for snacks, 10.8 percent for ice cream, and 13.7 percent for biscuits.

on September 20, 2011 in Korea's *Medical Today* documented that several retailers of products under this regulation delayed the reintroduction of MSRPs due to existing inventories, especially for items with shelf lives of about six months.

Second, producers decided not to reintroduce MSRPs for some products. At the government's request, producers declared the products for which they intended to reintroduce the price recommendations and their new levels, listed in Table A-1.[7] The fact that MSRPs were not reintroduced on some products helps us identify the effect of the reversal of the ban, which is described in more detail in the next section.

3 Data

The data comes from Nielsen Korea Retail Measurement Service and covers 253 processed food products from January 2010 to January 2012. The data contains monthly sales figures and average unit prices across five retailer categories for biscuits and pies, ramen, ice cream, and snacks, which are subject to the open price regulation, and products in cereal and yogurt, which are not. Cereal and yogurt products did not use MSRPs on their packaging either before or during the ban, nor after its reversal.

The sample consists of 24,837 product-retailer-month observations after dropping 4,563 observations due to missing prices or sales data. Table 1 presents descriptive statistics of the sample organized by product category. Price levels and dispersion for products subject to the open price regulation (Panel A) are comparable to those products not subject to it (Panel B). Sales-weighted average prices range from \$0.62 for ramen to \$13.27 for products in the biscuit and pie category. Products in the cereal and yogurt categories also have prices in this range: \$9.18 and \$5.35, respectively. Price dispersion is similar across categories. The coefficient of variation ranges from 22 to

[7] After the lift of the ban, MSRPs of most products were set at the same level as before the ban. Decreases in MSRPs were found for only 1 out of 12 products in the biscuit and pie category, 5 out of 19 products in the ice cream category, and 1 out of 11 products in the snack category. Only 1 product out of 19 in the ice cream category and 2 products out of 11 in the snack category increased their recommended prices. This information is not available for the ramen category.

57 percent for products under the policy and 22 to 61 percent for products not under the policy.

Table 2 describes the market structure for each product category. There are between two and five producers in each category, and prices within each category are similar across products and producers. The cereal and yogurt markets are more concentrated than categories affected by the open price regulation. In both categories two firms control more than 96 percent of the market, as compared to the market share of the two largest firms ranging from 62 to 86 percent for categories subject to the regulation. Also important to note is the fact that these manufacturers operate across multiple food categories, and in some cases the same manufacturer is subject to the regulation in one category but not another. For example, Lotte produces biscuits, ice cream, and snacks; Nongsim produces ramen, snacks, and cereal; and Bingrae produces yogurt and ice cream. We argue that the fact that cereal and yogurt belong to the processed foods industry and have similar prices and market structure, make them a suitable control group. In addition, using the MCIE press release of October 2011, we identify products under the regulation that included an MSRP after the ban was lifted and which products in the same categories did not reinstate MSRPs. These products are suitable to use as the control group to estimate the effect of the lift of the ban (see Table A-1 for a list of products).

We complement this data with the monthly producer price index (PPI) from the Bank of Korea for each of the 6 product categories. In order to control for retailer ownership structure we include an indicator variable equal to 1 if the retailer category is an independent store.

4 Empirical Analysis

In this section, we present the difference-in-differences methodology used to estimate the effect on prices of the MSRP ban and its reversal a year later. The econometric specification exploits the differential impact on products under the policy with comparable products that were not included in the regulation. We first present estimates of the effects of the ban and then estimate the impact on prices of the reverse regulation. We use various treatment and control group specifications.

The estimation of the two opposing policies allows us to address two concerns. First, any unobserved factor affecting the treatment or control group contemporaneously with the MSRP ban would not likely be present in the reverse direction after the lift of the ban. Second, we can test the validity of the estimated effect of MSRPs by comparing its sign and magnitude resulting from the ban and its subsequent reversal.

MSRP Ban

Let T_i denote an indicator variable which takes the value of 1 if a product, indexed by i, belongs to the treatment group and takes the value of 0 if it belongs to the control group. The treatment group includes products under the open price regulation policy: those in the biscuit and pie, ice cream, ramen, and snack categories. For this policy, the control group includes products in the cereal and yogurt categories. The treatment group indicator captures differences in the treatment and control group prior to the policy change. The specification includes an after-the-ban indicator, $AfterBan_{it}$, equal to 1 for the period after the ban and zero for the period before, and captures aggregate factors that affect prices prior to the implementation of the policy. In our case, we are interested in the effect of the policy on prices P_{it}. Our observations are for product i at time t and retailer category, however, for notational simplicity we omit the retailer index in the following difference-in-differences specification:

$$\ln P_{it} = \alpha + \beta AfterBan_{it} + \gamma T_i + \delta AfterBan_{it} \times T_i + \mathbf{x}_{it}\lambda + u_{it}.$$

The effect of the MSRP ban is captured by δ, the coefficient on the interaction variable which equals 1 for observations in the treatment group after the ban. The $\mathbf{x}_{it}$ vector includes an indicator for independent stores to capture differences in price strategies across retailers; PPI for each product category, which captures supply shocks from input prices; and a full set of controls for retailer, month, producer, product, and product category. Lastly, u_{it} are product-specific errors.

To alleviate concerns about the serial correlation of u_{it}, we follow Bertrand, Duflo, and Mullainathan (2004) by estimating the model with clustered robust standard errors, considering a product-retailer category as a cluster allowing an unrestricted variance-covariance structure (see also Donald and Lang, 2007).

Table 3 presents estimates of the overall effect of the ban of MSRPs on $\ln P_{it}$ for two different specifications. The first specification uses monthly observations for a window of time starting three months before and ending three months after the ban. The estimates indicate the ban led to a statistically significant, although moderate, 2.2 percent increase in prices. The second specification uses a window six months before and six months after the ban and yields similar results, that prices increased 2.3 percent after the ban. We obtain similar results in terms of magnitude and significance using bootstrapped standard errors.

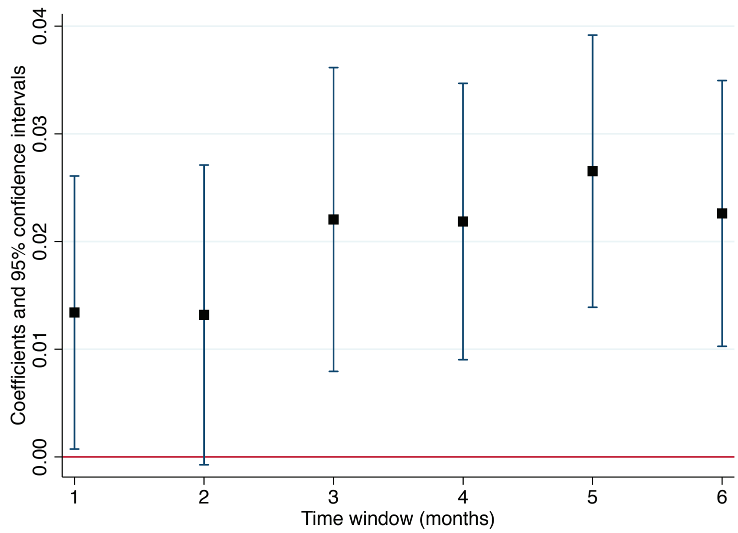


Figure 1: Coefficients of the Effect of the MSRP Ban

To test the robustness of the estimates, we use different time windows around the official date of the ban using the specification in Table 3. Figure 1 presents coefficients of the effect of the ban starting with a one-month window and increasing the window up to six months before and six months after the ban. For the one-month window the effect of the ban of prices is above 1 percent, but although the effect is of similar magnitude for the two-month window, it is not statistically significant at the 5 percent level. For longer time windows the effects on prices are larger than 2 percent and statistically significant.

In Table 4 we present estimates of the effect of the ban by product category. We use the same control group consisting of cereal and yogurt for all specifications. The effect of the ban is consistent across product categories, leading to price increases from 1.3 to 3.7 percent. For biscuits and pie, the estimates indicate that the ban increased prices 1.5 percent in both the three- and six-month windows. For ice cream, prices increase 2.5 and 3.7 percent, respectively, for the three- and six-month windows. For ramen, the effect is 1.3 percent in the three-month window, however the effect is not significant for the six-month window. Snacks experienced a price increase greater than 3 percent.[8]

These results are indicative of an overall price increase after the MSRP ban. One concern is the quality of the control group since cereal and yogurt products did not have an MSRP throughout the sample period. In the next section, we alleviate this concern by estimating the effect of the lift of the ban using a different control group, namely products that had MSRPs prior to the ban but had not reinstated them after the policy reversal.

MSRP Ban Reversal

In this section we present estimates of the lift of the MSRP ban. As mentioned above, this policy reversal was announced in June 2011, allowing price recommendations back on products as early as August 2011. The reinstatement of MSRPs was optional and not all manufacturers reintroduced MSRPs on their products. Consequently, products from the same food category are differentially subjected to the treatment of using an MSRP and we use those products that did not reinstate recommendations as a control group.

Table 5 presents the products that reinstated MSRPs after the reversal of the ban. This information was released by the MCIE on October 18, 2011, two and a half months after the

[8]Note that since there is no variation in PPI for snacks, it is dropped from the estimating equation. Similarly, there is no PPI variation for the three-month window for ramen.

ban was lifted, and indicates producers' intentions to include MSRPs on products. By this date, producers reinstated MSRPs on 12 out of 49 biscuit and pie products, 19 out of 94 ice creams, and 11 out of 40 snacks. The MCIE did not provide information on ramen products.[9]

Figure 2 shows the differential price change, over a time period 6 months prior to and 6 months after the ban reversal, of products that reinstated MSRPs compared to products that did not reinstate MSRPs. The figure illustrates that among products that did not reinstate MSRPs, prices increased from 1.4 to 4.6 percent across product categories, compared with products in the same categories that reinstated MSRPs after the reversal of the policy.

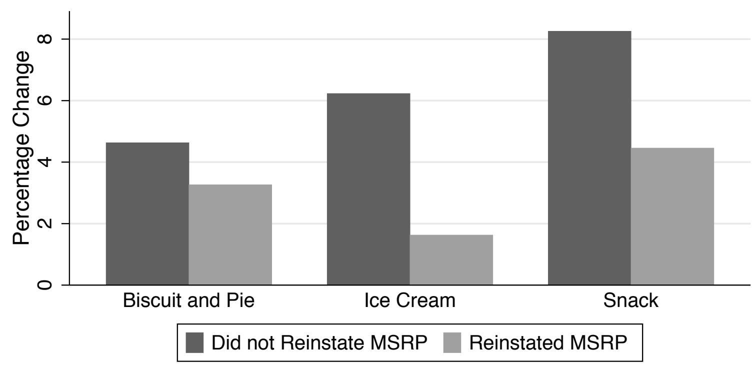


Figure 2: Average Price Change After the Reinstatement of MSRPs

There are two main distinctions from the econometric specification for the reversal of the ban compared with the specification for the ban. First, the treatment group includes only products that reinstated MSRPs, and the control group includes products covered by the ban that did not reintroduce MSRPs. Second, there is no homogeneous date when MSRPs were reintroduced for all products. We use an after reversal indicator variable that equals 0 before the reversal on August 2011, and equals 1 afterward. The effect of the MSRP ban reversal is captured by δ in the following

[9] As mentioned previously and illustrated in Table A-1 , the majority of MSRPs were set at the same levels as before the ban.

difference-in-difference specification:

$$\ln P_{it} = \alpha + \beta AfterReversal_{it} + \gamma T_i + \delta AfterReversal_{it} \times T_i + \mathbf{x}_{it}\lambda + u_{it}.$$

Table 6 presents the overall effect of the MSRP ban reversal on prices for our main specification. The effect of the policy for the three-month window is not statistically significant. This may be due to the fact that there was a short time between the announcement and the effective date of the lift of the ban and most producers' inventories, which did not include MSRPs, had not been exhausted. The effect for the six-month window is -2.6 percent, a similar magnitude but opposite sign of the estimated effect of the ban. We interpret the fact that the effect is absent in the first three months but present in the six-month specification as additional evidence that prices were affected only when MSRPs reappeared on product packaging.

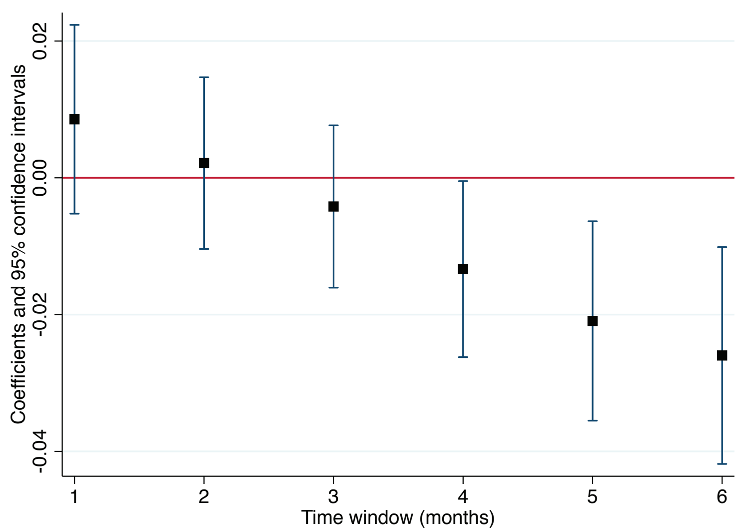


Figure 3: Coefficients of the Effect of the MSRP Ban Reversal

Figure 3 demonstrates the gradual effect of the lift of the ban by displaying estimated coefficients for different time windows around the official date of the MSRP ban reversal. There is a negative relationship between the time window and the policy coefficient. The effect of the ban reversal is not significant up to a three-month window and there is a statistically significant negative effect on prices for the four-month window specification which increases for the five-month and six-month windows.

In order to further test the robustness of our approach, Table 7 presents estimates of the policy reversal coefficient under several definitions of treatment and control groups. The first row is our main specification from Table 6, where ramen products are excluded from the treatment group and the control group includes products under the policy that did not reintroduce MSRP. In the second specification, we expand the control group to include cereal and yogurt products. The last two specifications include ramen products in the treatment group for two definitions of the control group—excluding and including cereal and yogurt in the control group, respectively. The effect of the policy across specifications is not significant in the three-month window, but it is negative and significant in the six-month window with a magnitude of -1.9 to -2.9 percent. Regressions for these specifications are provided in the Appendix tables A-2 to A-4.

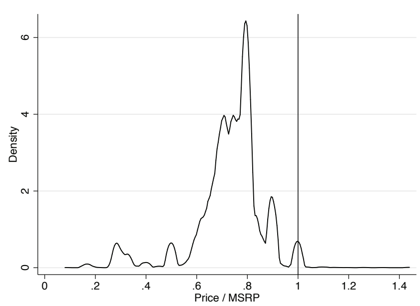
Table 8 presents estimates of the reversal of the policy for our main specification by product category. The policy coefficient is not significant for the three-month window for any product category, nor for the six-month window for biscuit and pie products. The effect for ice cream and snacks is -3 percent.

The empirical analysis in this section shows that the two opposite policies lead to two effects of opposite sign and similar magnitude. In general, this evidence indicates that the presence of MSRPs decreases prices.

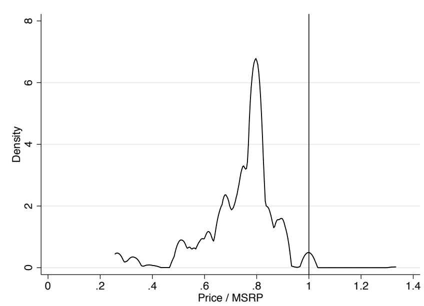
5 The Mechanism behind MSRPs

Having established empirically that the presence of MSRPs reduced prices, we investigate the mechanism behind this effect. We first conjecture that price recommendations act as price ceilings. For example, a manufacturer may punish retailers for exceeding his recommended price by withholding the product in the future or terminating the relationship altogether. Alternatively, the ceiling can be enforced by the behavior of consumers, who use the price recommendation as a reference point (Thaler, 1985; Puppe and Rosenkranz, 2011) and hesitate to pay prices above it.

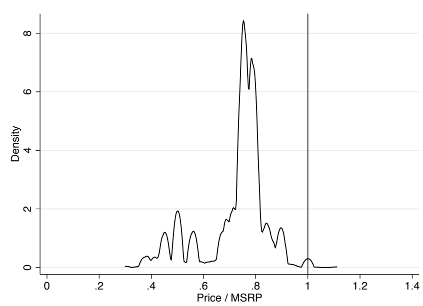
To find support for this theory we take a closer look at the distribution of prices. While the Nielsen data set used in the preceding analysis covers a comprehensive sample of products, it provides average sales-weighted prices across retailers and to examine whether MSRPs acted as price ceilings we are interested in the distribution of prices at individual stores. For this purpose, we use price data collected by the Korean Consumer Agency (KCA), consisting of weekly store-level prices for 16 products affected by the MSRP policy at 170 different stores. We use the KCA data to make some qualitative observations.



(a) Six Months Before the Ban



(b) Six Months After the Ban



(c) Six Months After the Ban Reversal

Figure 4: Relative Price Distributions of Products under the MSRP Regulation

Figure 4 depicts distributions of prices for products that used MSRPs prior to the ban, taken over three different one-month periods: six months prior to the ban, six months after the ban, and then six months after the reversal. The figures lead us to question whether MSRPs reduced prices by acting as binding ceilings. First, note that at each of the three times most prices are far below the MSRP, thus the recommended price does not bind as a constraint for a vast majority of retailers. Across the three snapshots, the proportion of prices more than 10 percent below MSRP is 96 percent prior to the ban, 97 percent during the ban, and 98 percent after the reversal. Second, there is no substantial increase in the number of stores charging prices above the recommendation after the ban, and no substantial decrease after the ban reversal. Thus, it seems that the MSRP does not bind even for retailers that set prices near the recommendation. We conclude that it is not likely that MSRPs acted as price ceilings.

An alternative explanation, proposed by a recent literature, argues that MSRPs are not explicit restraints but instead informative signals. In Buehler and Gartner (2012), MSRPs are used by a manufacturer to convey information about demand or cost to a retailer. The authors demonstrate that such communication is credible in an infinitely repeated game because there exists an equilibrium in which both firms have a shared objective of maximizing joint profits. In this environment, the manufacturer has no need to exert control over the retailer and instead uses the recommendation to share private information, thereby enabling the retailer to maximize joint profits more effectively. A ban of MSRPs could cut communication between the manufacturer and retailer and result in the retailer facing uncertainty over market conditions when setting prices, which may potentially lead either to higher or to lower prices on average. However, it is likely the manufacturer has means other than the MSRP to communicate with the retailer, thus in practice is not clear that a ban should have any observed effect.

Price recommendations may also be aimed at consumers. This is supported by the fact that MSRPs are often promoted through costly advertising and are visibly printed on product packaging. Lubensky (2011) proposes that MSRPs help consumers make search decisions by informing them of the distribution of retail prices, and demonstrates that providing this information is in the best interest of the manufacturer. Consumers then use the recommended price to determine which prices to accept and which prices to reject and continue to search. A ban on MSRPs results in consumers being uncertain about the distribution of retail prices that they face. Whether this uncertainty induces consumers to search more aggressively, which would intensify competition and decrease prices, or induces them to search less which would increase prices, is not immediately clear. To fix ideas and decompose the effects of uncertainty on consumer search, we present the following analysis.

MSRPs as Information for Consumers

A consumer with unit demand faces two sellers. The consumer first observes a price p at a randomly selected seller and then decides whether to visit the other seller. A visit costs $s > 0$ and reveals price p', which from the consumer's perspective is a random variable drawn from distribution $F(p, \theta)$. Having observed p', the consumer can accept it or return and accept p at no additional cost.

Parameter θ describes market conditions which affect the distribution of prices, for instance the realization of a shock to retailers' costs or aggregate demand. Let θ be uncertain with two equally likely outcomes, θ_l and θ_h. The regime with MSRPs is one in which the realization of θ is revealed to the consumer, and the regime without MSRPs is one in which the consumer remains uncertain about θ and consequently about the distribution of prices.

To describe the consumer's behavior in either regime, we first define the expected return to searching as

$$W(p,\theta) \equiv \int_{p'>p} (p - p') \ dF(p',\theta) \ = \ \int_{p'>p} F(p',\theta) \ dp'. \tag{1}$$

Search provides an option which is exercised whenever a lower price is observed, thus the value of searching depends on the price p at which the consumer starts. The latter equality follows from integration by parts.

In the environment with MSRPs in which the consumer observes θ, he optimally accepts at the first retailer if and only if the price does not exceed a threshold $p(\theta)$, which solves

$$s = W(p(\theta), \theta). \tag{2}$$

Let θ_l be the state associated with lower prices in the sense that $p(\theta_l) < p(\theta_h)$. In the environment without an MSRP the consumer's return to searching is uncertain. Given a threshold strategy $\tilde{p}$ is optimal,[10] it must satisfy

$$s = \mu(\tilde{p})W(\tilde{p}, \theta_h) + (1 - \mu(\tilde{p}))W(\tilde{p}, \theta_l), \tag{3}$$

in which $\mu(\tilde{p})$ is the consumer's posterior of the high price state θ_h. To get some intuition for how $\tilde{p}$ is determined, first conjecture that $\tilde{p} = p(\theta_l)$. Then the search decision is optimal when $\theta = \theta_l$ but the benefit to searching when $\theta = \theta_h$ is too low, and in this state the consumer is better off accepting some prices above $\tilde{p}$ instead of searching. Conversely, $\tilde{p} = p(\theta_h)$ yields the optimal search decision when $\theta = \theta_h$ but induces the consumer to accept prices he ought to reject when $\theta = \theta_l$. It follows that the threshold $\tilde{p}$ is in the interval $[p(\theta_l), p(\theta_h)]$ and optimally balances the loss from ex-post suboptimal search in both states. This idea is illustrated by re-arranging equation (3) as follows:

$$(1 - \mu(\tilde{p}))\Big(W(\tilde{p}, \theta_l) - W(p(\theta_l), \theta_l)\Big) = \mu(\tilde{p})\Big(W(p(\theta_h), \theta_h) - W(\tilde{p}, \theta_h)\Big).$$

Substituting in the definition of $W(p, \theta)$ from equation (1) obtains

$$(1 - \mu(\tilde{p})) \int_{p(\theta_l)}^{\tilde{p}} F(p', \theta_l)\ dp' = \mu(\tilde{p}) \int_{\tilde{p}}^{p(\theta_h)} F(p', \theta_h)\ dp'. \tag{4}$$

[10]Rothschild (1974) demonstrates that when the distribution of draws is uncertain, the consumer's optimal search policy is not necessarily a threshold strategy. We focus on price distributions that satisfy the necessary conditions for the existence of a threshold strategy, which requires that beliefs do not increase too quickly over a range of relevant prices.

The left-hand side describes the expected loss from searching too little when $\theta = \theta_l$ and the right-hand side describes the expected loss from searching too much when $\theta = \theta_h$.

Our main question is whether the consumer searches more or less, on average, when MSRPs are removed and he faces uncertainty. In other words, we compare the threshold $\tilde{p}$ used in the absence of MSRPs to the average of $p(\theta_l)$ and $p(\theta_h)$ when MSRPs are present. To make this comparison, we preform a decomposition of the effects of uncertainty, similar to that in Bénabou and Gertner (1993), into the relative losses from searching too little versus searching too much and the posterior likelihood that it is a state of low or high prices. We present the main ideas here and provide a more formal analysis and a discussion of how it relates to results in Bénabou and Gertner in the Appendix.

First, searching too little tends to be a costlier mistake than searching too much, due to the fact that search is an option. The ex-post loss from searching unsuccessfully is at most s, while the ex-post loss from not searching enough is potentially unbounded and increases as the search threshold increases. Formally this is reflected in equation (1), where the return to searching $W(p, \theta)$ is convex in threshold p. Thus, since $W_p(p, \theta)$ is increasing in p, the marginal loss at a threshold that is too high is greater than the marginal loss at a threshold that is too low.[11]

Second, the consumer's posterior $\mu(\tilde{p})$ can counteract the fact that searching too little is the costlier mistake. In absence of MSRPs, the consumer's prior that both θ_l and θ_h are equally likely is updated using Bayes' rule after observing the price at the first retailer. For example, if the first price p could only have occurred when $\theta = \theta_h$ then the consumer knows with certainty that his next price comes from distribution $F(p, \theta_h)$, and optimally searches if and only if $p \geq p(\theta_h)$. In general, consumers become pessimistic about the return to searching when the first observed price is associated more with θ_h and optimistic when it is associated more with θ_l. Having established that the threshold must come from $[p(\theta_l), p(\theta_h)]$, if prices in this interval are associated more with

[11] The logic applies in our setting with the caveat that ex-post search mistakes are made in two different distributions: too little search under $F(p, \theta_l)$ and too much search under $F(p, \theta_h)$. If the curvature of the distributions is sufficiently different, it is possible that the mistake of searching too much dominates. A more detailed discussion of this issue is presented in the Appendix.

state θ_h then uncertainty can reduce search by inducing a pessimistic posterior.

The removal of MSRPs thus has two potentially countervailing effects on search through the option value and the posterior, and either effect can dominate. The net of these two effects can be linked to an effect on prices through standard equilibrium logic. For instance, suppose that uncertainty increases search. This corresponds to an inward shift of demand, since consumers are more likely to reject any given price. The inward demand shift typically, though not necessarily,[12] then leads retailers to set lower prices. Thus one would expect that whenever uncertainty increases search, it also decreases prices.

In light of our main empirical finding that the presence of recommendations lowered prices, the theory of MSRPs as information to consumers is consistent if the posterior effect dominated the option value effect. In addition, in contrast to the theory of recommendations as explicit price restraints, this theory accommodates the evidence that most price changes occurred substantially below the MSRP level. Specifically, if the effect of MSRPs is to alter search thresholds then no explicit link between recommended prices and charged prices need exist.

6 Conclusion

In this paper, we contribute to the ongoing debate of whether manufacturers' attempts to exert influence over retail prices are pro- or anticompetitive. We exploit a natural experiment in which a ban of MSRPs was initially imposed and reversed a year later, and find that prices increased by 2.3 percent when MSRPs were banned and decreased by 2.6 percent after the ban's reversal.

Our empirical finding also sheds light on the mechanism by which MSRPs affect prices. Since recommendations lowered prices, we rule out theories that MSRPs are used to foster collusion or to maintain higher prices to ensure quality provision. By examining store-level prices, we also cast doubt on the theory that recommendations act as a form of maximum resale price maintenance.

[12]If the inward shift is accompanied also by a decrease in price elasticity, it is possible for profit-maximizing prices to rise. Moraga-Gonzáles, Sándor, and Wildenbeest (2013) demonstrate the possibility of such an effect in a search environment.

Instead, we propose an alternative explanation in which MSRPs provide information to searching consumers.

An important lesson from this analysis is the need to understand the manufacturer's motives in exerting influence over retail prices and the mechanism by which the manufacturer's actions affect market outcomes. The Korean ban was intended to promote competition and reduce prices but had the opposite effect, which led to its reversal a year later. While our empirical work shows that in this market MSRPs reduced prices, it does not imply that MSRPs are in general pro-competitive. Given that it is unlikely that MSRPs acted as an explicit restraint and that instead their role was to provide information, our theoretical framework demonstrates that the price effect of MSRPs is ambiguous and may differ across markets. Effective regulation of manufacturer price recommendations should thus treat industries with different characteristics on a case by case basis.

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Appendix 1. Analysis of MSRPs as Information

We investigate the characteristics of the optimal search threshold under uncertainty which solves equation (3). We demonstrate that the mistake of searching too little tends to be, but does not have to be larger in magnitude than the mistake of searching too much. We then show that the posterior effect can dominate.

As a benchmark, consider first an example in which θ shifts the distribution of prices but does not change its curvature, as in Figure 5. Namely, there is an underlying distribution $G(p)$ and for every θ the distribution of prices is given by $F(p, \theta) = G(p - \theta)$. If θ is known then the optimal threshold solves

$$s = \int_{\underline{p}+\theta}^{p(\theta)} G(p' - \theta)\ dp' \ = \ \int_{\underline{p}}^{p(\theta)-\theta} G(p)\ dp.$$

Note that in this formulation, $p(\theta) - \theta$ is constant with respect to θ, thus $F(p(\theta), \theta) = G(p(\theta) - \theta))$ is also constant with respect to θ. Put differently, the same percentile price is used as the threshold regardless of state θ.

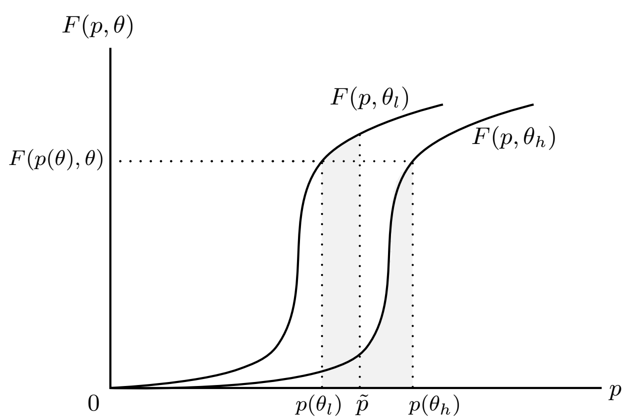

Figure 5: θ changes the mean of the price distribution

Figure 5 shows two price distributions, one for each state θ_l and θ_h, with the latter resulting from a horizontal rightward shift. The shaded area for prices between $p(\theta_l)$ and $\tilde{p}$ represents the loss from searching too little when $\theta = \theta_l$, as described in equation (3). Similarly, the shaded area between prices $\tilde{p}$ and $p(\theta_h)$ represents the losses from searching too much when $\theta = \theta_h$. Importantly, due to the nature of the shift, $F(p(\theta_l), \theta_l) = F(p(\theta_h), \theta_h)$. This implies that at any $p \in [p(\theta_l), p(\theta_h)]$, the value of searching is strictly more sensitive in state θ_l than in state θ_h, as evidenced in the figure by the fact that $F(p, \theta_l) > F(p, \theta_h)$ for prices in this range. As a result, the loss from searching too little dominates the loss from searching too much, and tends to reduce the threshold $\tilde{p}$ toward $p(\theta_l)$.

The discussion above outlines the tendency for losses from searching too little to be larger than those from searching too much. However in doing so we abstracted from the role of beliefs $\mu(\tilde{p})$, as given in equation (3). The following lemma incorporates the belief to formalize this idea.

Lemma 1 *There exists* $\mu > \frac{1}{2}$ *such that if* $\mu(\tilde{p}) \in [0, \mu]$ *then* $\tilde{p} \leq \frac{1}{2}p(\theta_l) + \frac{1}{2}p(\theta_h)$.

Proof First note the following two inequalities:

$$\begin{aligned} \frac{1}{2}(\tilde{p} - p(\theta_l))F(p(\theta_l), \theta_l) \ &\leq \ (1 - \mu(\tilde{p})) \int_{p(\theta_l)}^{\tilde{p}} F(p', \theta_l) \ dp' \\ \frac{1}{2}(p(\theta_h) - \tilde{p})F(p(\theta_h), \theta_h) \ &\geq \ \mu(\tilde{p}) \int_{\tilde{p}}^{p(\theta_h)} F(p', \theta_h) \ dp' \end{aligned}$$

Both inequalities are strict for $\mu(\tilde{p}) = \frac{1}{2}$, thus the inequalities would still hold for a slightly larger $\mu(\tilde{p})$. By definition of $\tilde{p}$ in equation (3), the right-hand sides of both inequalities are equal, thus it follows that

$$\frac{1}{2}(\tilde{p} - p(\theta_l))F(p(\theta_l), \theta_l) \ \leq \ \frac{1}{2}(p(\theta_h) - \tilde{p})F(p(\theta_h), \theta_h).$$

Using the fact that in this example, $F(p(\theta_l), \theta_l) = F(p(\theta_h), \theta_h)$, we obtain that

$$\tilde{p} \leq \frac{1}{2}(p(\theta_l) + p(\theta_h)). \quad \blacksquare$$

Lemma 1 and the corresponding figure show that because searching too little is a bigger mistake than searching too much, uncertainty increases search on average. This follows from the fact that for any price $p \in [p(\theta_l), p(\theta_h)]$, the marginal value of search $F(p, \theta)$ is higher when $\theta = \theta_l$ than when $\theta = \theta_h$. However if θ is not simply a horizontal shifter, it is possible for the marginal value of search to be higher when $\theta = \theta_h$, inducing $\tilde{p} > \frac{1}{2}p(\theta_l) + \frac{1}{2}p(\theta_h)$. For instance, imagine that in state θ_l the price distribution has a long left tail relative to the distribution in state θ_h, as depicted to scale in Figure 6. Now, for any $\tilde{p} \in [p(\theta_l), p(\theta_h)]$, the value of searching is more sensitive to the threshold in state θ_h than it is in state θ_l, which causes the mistake in state θ_h to be of a higher magnitude. The $\tilde{p}$ in Figure 6, which is closer to $p(\theta_h)$ than to $p(\theta_l)$, is chosen to equalize the two shaded regions, which corresponds to the optimal threshold given the prior belief $\mu(\tilde{p}) = \frac{1}{2}$.

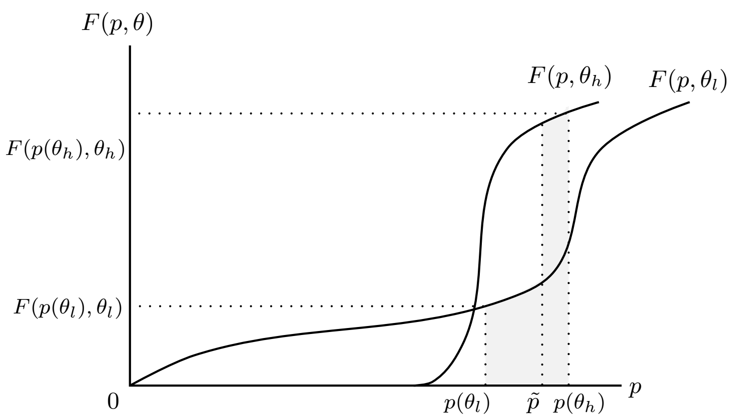

Figure 6: θ changes the mean and curvature of the price distribution

Lemma 2 *There exists $\mu < \frac{1}{2}$ such that for any $\mu(\tilde{p}) \in [\mu, 1]$, if $F(p(\theta_h)\ ,\theta_l) < F(p(\theta_l), \theta_h)$ then $\tilde{p} > \frac{1}{2}p(\theta_l) + \frac{1}{2}p(\theta_h)$.*

Proof Once again the proof follows from a sequence of inequalities using equation (3).

$$
\begin{aligned}
\frac{1}{2}(\tilde{p} - p(\theta_l))F(p(\theta_h), \theta_l) &\geq (1 - \mu(\tilde{p})) \int_{p(\theta_l)}^{\tilde{p}} F(p', \theta_l) dp' \\
&= \mu(\tilde{p}) \int_{\tilde{p}}^{p(\theta_h} F(p', \theta_h) dp' \\
&\geq \frac{1}{2}(p(\theta_h) - \tilde{p})F(p(\theta_l), \theta_h)
\end{aligned}
$$

Both inequalities are strict whenever $\mu(\tilde{p}) = \frac{1}{2}$, thus the inequalities continue to hold for slightly lower $\mu(\tilde{p})$. Comparing the first and last terms obtains

$$
\begin{aligned}
\tilde{p} - p(\theta_l) &\geq p(\theta_h) - \tilde{p} \\
\tilde{p} &\geq \frac{1}{2}(p(\theta_l) + p(\theta_h)). \quad \blacksquare
\end{aligned}
$$

To summarize, the ex-post mistake in state θ_l is to accept prices above the full-information threshold that one should reject and the ex-post mistake in state θ_h is to reject prices below the full-information threshold that one should accept. If the full information thresholds in both states are associated with the same probability of observing a lower price, the mistake of searching too little is of a higher magnitude. However, if at the full information threshold the probability of a lower price is smaller in state θ_l, then the mistake of searching too much might be the costlier of the two. This can occur, for instance, if state θ_l is associated with a small possibility of very low prices. Thus, the relative magnitudes of the two mistakes in Equation (3) are generally ambiguous and depend on particular characteristics of the price distributions.

Aside from the losses associated with suboptimal search in each state, the other determinant of $\tilde{p}$ is the posterior $\mu(\tilde{p})$, which describes the relative likelihood of each loss. The posterior is formed using Bayes rule and, without additional structure on the distribution of prices, it is difficult to ascertain whether the posterior induces more or less search. In the preceding example in which

$F(p, \theta) = G(p - \theta)$, the posterior at any price $p \in [p(\theta_l), p(\theta_h)]$ is

$$\mu(p) = \frac{g(p - \theta_h)}{g(p - \theta_h) + g(p - \theta_l)}.$$

If $g(p)$ is uniform, the posterior $\mu(p) = \frac{1}{2}$ equals the prior and, as demonstrated in Lemma 1, there will be more search under uncertainty. However, if $g(p)$ is decreasing in this range, then the posterior becomes $\mu(p) > \frac{1}{2}$, inducing consumers to be more pessimistic about the return to searching and potentially to search less under uncertainty.

The uncertainty ensuing from the removal of MSRPs thus has two potentially countervailing effects on search through the option value and the posterior. When price distributions $F(p, \theta)$ are specified exogenously, demonstrating that the posterior effect may dominate the option value effect is trivial. However, it needs to be shown that the posterior effect may dominate the option value effect when attention is restricted to equilibrium price distributions. That is, in an environment in which prices and search decisions are chosen simultaneously.

Obtaining such a result in a general setting remains an open problem. However, Bénabou and Gertner (1993) provides two examples, one analytic and one numerical, in which the posterior effect can dominate. In their model, the consumer faces the same search decision as in our model when price distributions are endogenous, with the two firms simultaneously setting prices after observing the realizations of a private and an aggregate cost shock. The authors find similar countervailing effects as we identify above, and demonstrate that an increase in uncertainty can lead to less search and higher prices.

While Bénabou and Gertner find that removing uncertainty can increase search in equilibrium, we need to interpret their result carefully with respect to our application to MSRPs. Specifically, the authors examine the effect on uncertainty on search by reducing the variance of the aggregate shock, and compare the search threshold $\tilde{p}$ when state θ is uncertain to the search threshold $p\left(\frac{1}{2}\theta_l + \frac{1}{2}\theta_h\right)$ when the average state is realized with certainty. An MSRP however does not remove fluctuations in aggregate conditions, rather it removes the information problem. The appropriate comparison then is that of $\tilde{p}$ with the average of the full information thresholds $\frac{1}{2}p(\theta_l) + \frac{1}{2}p(\theta_h)$. In principle, these two comparisons may differ from one another depending on the curvature of $p(\theta)$.

Table 1: Descriptive Statistics

Product Category	Number of Products	Number of Producers	Sales (M$)	Unit Price ($) Avg.	Unit Price ($) Std. Dev.	Obs.
Panel A. Products under Policy						
Biscuit and Pie	49	3	988	10.53	5.97	5,562
Ice Cream	94	3	2,776	5.06	1.88	7,697
Ramen	45	4	1,609	0.62	0.14	4,668
Snacks	40	5	850	13.27	4.97	4,466
Panel B. Products not under Policy						
Cereal	21	2	303	9.18	2.04	2,023
Yogurt	4	4	210	5.35	3.27	421

Note: The table presents descriptive statistics for different grocery products under the MSRP policy change and products not subject to the policy. The prices presented are weighted by sales. The unit of measurement is kilograms for biscuits, pie, snacks and cereal and liters for ice cream and yogurt. We use the average exchange rate in 2010 and 2011 of 1,132 Won/USD. Source: Nielsen Korean Retail Measurement Service.

Table 2: Market Shares and Prices by Producer

Firm	Market Share	Unit Price ($) Avg.	Unit Price ($) Std. Dev.
Panel A. Products under Policy			
Biscuits and Pie (Sales =M$988)			
Lotte	42.2	9.82	3.74
Haetae	32.1	12.59	8.73
Orion	25.7	9.12	3.54
Ice Cream (Sales=M$2,776)			
Lotte	46.6	5.15	1.67
Bingrae	32.4	4.78	2.13
Haetae	21.1	5.30	1.84
Ramen (Sales=M$1,609)			
Nongsim	75.0	0.62	0.13
Oddugi	11.2	0.58	0.15
Paldo	8.0	0.69	0.12
Samyang	5.8	0.63	0.14
Snack (Sales=M$850)			
Orion	37.6	14.91	3.50
Nongsim	24.6	10.98	4.35
Lotte	19.1	13.82	7.05
Haetae	16.3	13.05	4.14
Samyang	2.4	8.16	2.42
Panel B. Products not under Policy			
Cereal (Sales=M$303)			
Dongseo	51.1	9.63	2.02
Nongsim	48.9	8.71	1.96
Yogurt (Sales=M$210)			
Bingrae	63.4	4.89	0.54
Namyang	32.9	6.16	5.55
Maeil	2.9	6.18	0.26
Yakult	0.8	5.54	0.26

Note: We use the average exchange rate in 2010 and 2011 of 1,132 Won/USD.

Table 3: Estimates of the Effect of the MSRP Ban

	3-month Window		6-month Window	
Variable	Coeff.	Std. Err.	Coeff.	Std. Err.
After Ban x Treated	0.022	(0.007)***	0.023	(0.006)***
After Ban	-0.019	(0.007)***	-0.004	(0.006)
Treated	0.802	(0.063)***	1.381	(0.064)***
lnPPI	0.238	(0.125)*	0.375	(0.128)***
Independent Store	-0.398	(0.018)***	-0.036	(0.009)***
Constant	7.814	(0.593)***	7.162	(0.605)***
Fixed Effects:				
Retailer Category	Y		Y	
Month	Y		Y	
Producer	Y		Y	
Product Category	Y		Y	
Product	Y		Y	
R-squared	0.990		0.987	
Observations	5,726		11,351	

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. The sample includes products under the MSRP ban as the treatment group and cereal and yogurt products as the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table 4: Effect of the MSRP Ban by Product Category

	Biscuits and Pie		Ice Cream		Ramen		Snack	
	3-month	6-month	3-month	6-month	3-month	6-month	3-month	6-month
Variable	Coeff.	Coeff.	Coeff.	Coeff.	Coeff.	Coeff.	Coeff.	Coeff.
After Ban x Treated	0.015	0.015	0.025	0.037	0.013	0.007	0.032	0.034
	(0.007)**	(0.006)**	(0.010)**	(0.011)***	(0.007)**	(0.006)	(0.010)***	(0.010)***
After Ban	-0.018	-0.003	0.005	0.008	-0.015	0.003	-0.007	-0.002
	(0.008)**	(0.008)	(0.008)	(0.008)	(0.008)*	(0.008)	(0.008)	(0.011)
Treated	0.295	0.512	0.447	0.382	-3.012	-2.726	-0.381	0.181
	(0.062)***	(0.057)***	(0.047)***	(0.044)***	(0.023)***	(0.023)***	(0.021)***	(0.017)***
$\ln PPI$	0.090	-0.185	0.250	0.359		0.707		
	(0.111)	(0.068)***	(0.152)	(0.144)**		(0.175)***		
Independent Store	-0.021	-0.026	-0.163	-0.155	-0.083	-0.146	-0.114	-0.114
	(0.012)*	(0.011)**	(0.022)***	(0.020)***	(0.014)***	(0.014)***	(0.009)***	(0.010)***
Constant	8.539	9.822	7.787	7.265	8.934	5.599	8.935	8.923
	(0.529)***	(0.325)***	(0.718)***	(0.682)***	(0.028)***	(0.825)***	(0.050)***	(0.046)***
Fixed Effects:								
Retailer Cat.	Y	Y	Y	Y	Y	Y	Y	Y
Month	Y	Y	Y	Y	Y	Y	Y	Y
Producer	Y	Y	Y	Y	Y	Y	Y	Y
Product	Y	Y	Y	Y	Y	Y	Y	Y
R-squared	0.960	0.957	0.911	0.892	0.998	0.997	0.948	0.930
Observations	1,849	3,682	2,379	4,675	1,592	3,177	1,598	3,177

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. Products under the MSRP ban in the sample are sorted into 4 categories—biscuits and pie, ice cream, ramen, and snacks —and each works as the treatment group while cereal and yogurt products serve as the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table 5: Decision to Include MSRPs after Lift of the Ban

Product Category	MSRP	No MSRP	Unknown	Total
Biscuits and Pie	12	37	0	49
Ice Cream	19	75	0	94
Ramen	—	—	46	46
Snacks	11	22	7	40

Note: The table presents the number of products by category that included a MSRP after the lift of ban on MSRP. Source: Korean Ministry of Commerce, Industry, and Energy, press release, Oct. 18, 2011. This information was only available for biscuits and pie, ice cream, and snack product categories.

Table 6: Estimates of the Effect of the MSRP Ban Reversal
Control Group: Products That Did Not Include MSRP

	3-month Window		6-month Window	
Regressor	Coeff.	Std. Err.	Coeff.	Std. Err.
After Reversal x Treated	-0.004	(0.006)	-0.026	(0.008)***
After Reversal	0.022	(0.008)***	0.093	(0.011)***
Treated	0.268	(0.014)***	0.325	(0.036)***
lnPPI	0.196	(0.080)**	0.207	(0.071)***
Independent Store	-0.156	(0.010)***	-0.441	(0.015)***
Constant	8.991	(0.397)***	8.841	(0.353)***
Fixed Effects:				
Retailer Category	Y		Y	
Month	Y		Y	
Producer	Y		Y	
Product Category	Y		Y	
Product	Y		Y	
R-squared	0.961		0.95	
Observations	4,216		8,309	

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. The sample includes products under the MSRP ban. Products which printed MSRP after the lift of the ban are in the treatment group, while products which did not print MSRP after the lift of the ban are in the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table 7: Robustness Check of Policy Effect of Reinstating MSRP

Treatment includes Ramen	Control includes Cereal & Yogurt	3-month Window		6-month Window	
		Coeff.	Std. Err.	Coeff.	Std. Err.
N	N	-0.004	(0.006)	-0.026	(0.008)***
N	Y	0.000	(0.006)	-0.019	(0.008)**
Y	N	-0.007	(0.005)	-0.029	(0.007)***
Y	Y	-0.002	(0.005)	-0.020	(0.006)***

Note: The table presents policy coefficients from difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. The sample includes products under the MSRP ban. Products which include MSRP after the lift of the ban are in the treatment group. Since ramen products cannot be identified as including MSRP, ramen products were excluded from the estimation in the first two specifications and included in the last two. Products which did not include MSRP after the lift of the ban are the control group. The control group is extended to include cereal and yogurt products as indicated. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table 8: Estimates of the Effect of Reinstating MSRP by Product Category
Control Group: Products That Did Not Include MSRP

	Biscuits and Pie		Ice Cream		Snack	
	3-month	6-month	3-month	6-month	3-month	6-month
Variable	Coeff.	Coeff.	Coeff.	Coeff.	Coeff.	Coeff.
After Reversal	0.007	-0.017	-0.014	-0.030	0.000	-0.030
x Treated	(0.012)	(0.012)	(0.010)	(0.015)**	(0.009)	(0.010)***
After Reversal	0.021	0.056	0.017	0.034	0.041	0.085
	(0.013)*	(0.017)***	(0.012)	(0.029)	(0.006)***	(0.007)***
Treated	-0.810	-0.763	0.016	0.303	-0.069	-0.150
	(0.050)***	(0.070)***	(0.063)	(0.044)***	(0.028)**	(0.029)***
lnPPI	0.176	0.291	-0.195	1.082		
	(0.130)	(0.156)*	(0.531)	(0.347)***		
Independent Store	-0.171	-0.160	-0.452	-0.449	-0.115	-0.012
	(0.010)***	(0.009)***	(0.026)***	(0.023)***	(0.010)***	(0.010)
Constant	9.573	8.939	10.562	3.372	9.694	9.743
	(0.658)***	(0.789)***	(2.728)***	(1.807)*	(0.034)***	(0.032)***
Fixed Effects:						
Retailer Category	Y	Y	Y	Y	Y	Y
Month	Y	Y	Y	Y	Y	Y
Producer	Y	Y	Y	Y	Y	Y
Product	Y	Y	Y	Y	Y	Y
R-squared	0.957	0.945	0.870	0.834	0.947	0.948
Observations	1,378	2,734	1,923	3,760	915	1,815

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. Products under the MSRP ban in the sample are sorted into 3 categories—biscuits and pie, ice cream, and snacks. For each category, products which printed MSRP after the lift of the ban are in the treatment group, while products which did not print MSRP after the lift of the ban are in the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table A-1: MSRP Before the Ban and After the Lift of the Ban

Biscuits and Pie			Ice Cream			Snack		
Producer & Product	MSRP (Won) Before	After	Producer & Product	MSRP (Won) Before	After	Producer & Product	MSRP (Won) Before	After
Haetae			*Bingrae*			*Haetae*		
ace	1,000	1,000	bbangdoah	1,500	1,500	matdongsan	1,200	1,200
chocopick	1,500	1,500	chiness	1,000	900	osaeckgamja	1,500	1,500
ddangkongrae	3,600	3,400	metacorn	1,500	1,500	ossazze	1,200	1,000
gaerangwaja	1,000	1,000	nuribar	1,000	900			
sarubia	1,200	1,200	ssamanko	1,500	1,500	*Lotte*		
			thewiisanang	1,000	1,000	oing	1,200	1,200
Lotte						sunchip	1,200	1,200
jeck	1,000	1,000	*Haetae*					
lottechocopie	3,000	3,000	babambar	1,000	1,000	*Nongsim*		
			bravocorn	1,500	1,500	bananakick	700	800
Orion			hodumaru	1,000	1,000	saewookkang	800	900
chocopie	3,200	3,200	nugabar	1,000	1,000			
goraebab	700	700	ssangssangbar	1,000	1,000	*Orion*		
mizzblack	700	700				daedannacho	1,200	1,200
ott	5,000	5,000	*Lotte*			dodonacho	1,200	1,200
wehas	700	700	bingbingbar	1,000	900	ogamja	1,200	1,200
			jawsbar	1,000	1,000	orionsun	1,200	1,200
			nukebar	1,000	900			
			screwbar	1,000	1,000			
			seolraeim	1,500	1,600			
			tornado	1,000	900			
			wangsubak	1,000	1,000			
			worldcorn	1,500	1,500			

Source: Korean Ministry of Commerce, Industry, and Energy, press release, Oct 18, 2011. This information was only available for biscuits and pie, ice cream, and snack product categories.

Table A-2: Estimates of the Effect of Reinstating MSRP
Treatment Group: Excludes Ramen
Control Group: Products with no MSRP and Cereal & Yogurt

	3-month Window		6-month Window	
Regressor	Coeff.	Std. Err.	Coeff.	Std. Err.
After Reversal x Treated	0.000	(0.006)	-0.019	(0.008)**
After Reversal	0.014	(0.007)**	0.065	(0.008)***
Treated	0.263	(0.015)***	0.237	(0.035)***
lnPPI	0.302	(0.076)***	0.400	(0.063)***
Independent Store	-0.435	(0.017)***	-0.042	(0.009)***
Constant	9.101	(0.400)***	7.975	(0.313)***
Fixed Effects:				
Retailer Category	Y		Y	
Month	Y		Y	
Producer	Y		Y	
Product Category	Y		Y	
Product	Y		Y	
R-squared	0.956		0.948	
Observations	4,826		9,504	

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. The sample includes products under the MSRP ban. Products which printed MSRP after the lift of the ban are in the treatment group, while products which did not print MSRP after the lift of the ban as well as cereal and yogurt products are in the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table A-3: Estimates of the Effect of Reinstating MSRP
Treatment Group: Includes Ramen
Control Group: Products with No MSRP

	3-month Window		6-month Window	
Regressor	Coeff.	Std. Err.	Coeff.	Std. Err.
After Reversal x Treated	-0.007	(0.005)	-0.029	(0.007)***
After Reversal	0.019	(0.007)***	0.075	(0.009)***
Treated	0.198	(0.016)***	0.321	(0.035)***
lnPPI	0.298	(0.068)***	0.387	(0.056)***
Independent Store	-0.057	(0.009)***	-0.435	(0.014)***
Constant	9.189	(0.366)***	7.959	(0.281)***
Fixed Effects:				
Retailer Category	Y		Y	
Month	Y		Y	
Producer	Y		Y	
Product Category	Y		Y	
Product	Y		Y	
R-squared	0.992		0.990	
Observations	5,545		10,927	

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. The sample includes products under the MSRP ban. Products which printed MSRP after the lift of the ban are in the treatment group, while products which did not print MSRP after the lift of the ban are in the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).

Table A-4: Estimates of the Effect of Reinstating MSRP
Treatment Group: Includes Ramen
Control Group: Products with No MSRP and Cereal & Yogurt

	3-month Window		6-month Window	
Regressor	Coeff.	Std. Err.	Coeff.	Std. Err.
After Reversal x Treated	-0.002	(0.005)	-0.020	(0.006)***
After Reversal	0.016	(0.005)***	0.061	(0.007)***
Treated	0.288	(0.017)***	0.312	(0.036)***
lnPPI	0.371	(0.066)***	0.501	(0.055)***
Independent Store	-0.064	(0.009)***	-0.434	(0.014)***
Constant	8.734	(0.360)***	7.400	(0.275)***
Fixed Effects:				
Retailer Category	Y		Y	
Month	Y		Y	
Producer	Y		Y	
Product Category	Y		Y	
Product	Y		Y	
R-squared	0.991		0.990	
Observations	6,155		12,122	

Note: The table presents difference-in-differences estimates using $\ln(P_{it})$ as the dependent variable. The sample includes products under the MSRP ban. Products which printed MSRP after the lift of the ban are in the treatment group, while products which did not print MSRP after the lift of the ban as well as cereal and yogurt products are in the control group. Clustered robust standard errors are in parentheses. The notation *** indicates significant at 1% level, ** at 5% level, * at 10% level. A three-month window denotes observations from 3 months before and 3 months after the lift of the ban (similarly for the six-month window).