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A Meta-Analysis of National Brand and Store Brand Cross-Promotional Price Elasticities

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Abstract

This paper investigates whether price discounts by national brands influence private-label sales and vice versa through meta-analysis of 261 cross-price elasticity estimates from sixteen product-chains. On average, price reductions by national brands and private labels have more or less equal influence on each others' sales. However, there is greater variation in the effect of private-label price cuts across national brands. National brands with large market shares decrease private-label sales through price cuts but are seldom affected by private-label discounts. National brands with lower relative price have greater influence on private-label sales and are also affected more by private-label price cuts.

Private labels or store brands are generally brands owned, controlled, and sold exclusively by the retailers. Private labels in grocery products account for over $48 billion and have been growing rapidly (Hoch and Banerji, 1993). According to Information Resources, Inc., private labels bagged 19.7 percent of supermarket unit sales in 1993 compared to 15.8 percent in 1988 (Miller, 1995). One important basis for selling private labels is the price differential between store brands and national brands. Recent cross-category studies (McMaster, 1987; Sethuraman, 1992; Hoch and Banerji, 1992) have found a negative relationship between price differential and private-label share across categories—that is, the higher the price differential between national brand and store brand in a category, the *lower* is the market share. This result has been picked up by the popular press (see Gibson, 1992) and interpreted to imply that price differential is not an important determinant of private label share. Based on this finding and other considerations, some researchers (Hoch and Banerji, 1993; Sethuraman, 1992) have advocated that national brand manufacturers should perhaps focus less on price reduction and more on other aspects, such as product quality.

In response to these findings, Raju, Sethuraman, and Dhar (1995) have analytically demonstrated that cross-category studies may not reflect the "true" effect of price on private-label sales. An appropriate method for assessing the price effects would be to analyze within-category data. Relatedly, Blattberg and Wisniewski (1989) theorize and show evidence in four products from one chain that, when the higher-price-tier, higher-quality (national) brands are price promoted, they draw sales from their own price-tier competitors and from the tier below (private brands). On the other hand, if the lower-quality, lower-price-tier (private) brands are promoted, they rarely take sales from the (national brand) tier above.

Thus, as national brand manufacturers and retailers engage in a price battle, the following questions arise:

- Overall, do national brand price reductions have little effect on private-label sales as alluded to in Hoch and Banerji (1993) and as reported in some business press?
- Or, overall, do national brand price reductions "hurt" (significantly decrease) private-label sales, while private-label price reductions have little effect on national brand sales, as theorized by Blattberg and Wisniewski (1989)?
- What types of national brands hurt (significantly decrease) private-label sales through their price cuts?
- What types of national brands are hurt by private-label price cuts?

This paper addresses these questions by studying the effect of short-term price changes (price discounts) of national brands (private labels) on the sales of private labels (national brands) across several products and stores. In particular, 261 cross-price elasticity estimates across sixteen product-chain combinations covering seventy-one brand observations are meta-analyzed to provide some limited generalizations in a manner similar in spirit to the studies by Assmus, Farley, and Lehmann (1984), Bolton (1989), and Leone and Schultz (1980).

The paper is organized as follows. First, I describe the data and the method used for estimating the cross-price elasticites. Then, I present the meta-analysis of these cross-price elasticities. Finally, I discuss the key empirical results and provide the limitations and directions for future research.

1. Data and estimation of cross-price elasticities

1.1. Data

The data are store-level supermarket scanner data obtained from Information Resources, Inc. (IRI). The empirical analysis is performed on six product categories: (1) four-roll white bathroom tissue, (2) forty-count fabric softener sheets, (3) 5 lb. all-purpose flour, (4) 16 oz. margarine, (5) 64–95.9 oz. orange juice from concentrate, (6) half-size water-based canned tuna. Data for these six product categories are obtained from three stores in different locations belonging to three different chains. However, data are unavailable for flour and orange juice in one of the chains. Thus, there are sixteen data sets or product-chains comprising seventy-three brands (fifty-six national brands, sixteen private-label brands, and one generic brand). Each data set contains weekly information on unit sales by item, price by item, whether the item was discounted and the deal percentage, as well as display and feature information. Data are available for 104 weeks during 1991–1993. Details of aggregate frequency of discounts, size of price cuts, and the percentage sold on deal are provided in Table 1 for each of the sixteen product-chains. In all the data sets, there appear to be substantial price discount activity though the extent of such activity varies by product and store.

Table 1. Aggregate discount data and cross-price elasticities.

Product	Chain	Frequency of Discount (%)[a]	Average % Discount[b]	% Sold on Deal[c]	Average Cross-Price Elasticity
Bathroom tissue	A	44.3	14.8	70.3	.64
	B	39.1	21.8	47.4	.38
	C	42.8	21.7	58.7	.31
Fabric softener	A	36.3	10.2	38.3	.30
	B	25.6	13.8	19.8	.59
	C	22.5	13.4	24.3	.87
Flour	A	44.3	17.2	39.0	.70
	C	38.9	23.7	37.6	.48
Margarine	A	44.8	21.4	42.5	.38
	B	42.5	27.2	28.6	.24
	C	48.1	26.8	64.2	.66
Orange juice	A	44.9	21.4	48.0	.52
	B	26.9	14.6	71.9	.84
Tuna	A	43.8	23.9	74.7	.72
	B	39.2	22.0	65.6	.44
	C	29.1	18.7	43.0	.92
Total		38.3	19.5	48.4	.56

a. $\dfrac{\text{Number of weeks on discount}}{\text{Total number of weeks}} \times 100.$

b. $\dfrac{\text{Regular price-discounted price}}{\text{Regular price}} \times 100.$

c. $\dfrac{\text{Actual sales} - \text{Base (expected regular price) sales}}{\text{Actual sales}} \times 100.$

1.2. *Estimation procedure*

Of the seventy-three brands, one national brand is unavailable during part of the analysis period and one (generic) brand has low price variation and poor model fit. They are excluded and the remaining seventy-one brands (fifty-five national brands and sixteen private labels) are used in the analysis. I estimate each of the seventy-one brand sales models separately and obtain measures of own- and cross-price elasticities.

All 104 weekly observations are used for the estimation (no holdout sample). The econometric model relates the weekly sales of each brand to own price, competitors' price, and other variables (covariates) that might affect brand sales. The covariates are display and feature indicators for the estimation brand as well as competitors' brands. The display (feature) indicator for a brand takes a value 1 if the brand is dispalyed (featured) during that week and 0 otherwise. I also include a seasonality indicator for flour and margarine that takes a value 1 during November-December holiday season and 0 otherwise.

Three commonly used functional forms (Blattberg and Wisniewski, 1989; Bolton, 1989) are estimated for each brand: (1) a linear model where the dependent variable is unit sales and the independent price variables are unit prices, (2) a semi-log model where the dependent variable is logarithm of unit sales and the independent price variables are unit prices, and (3) a double-log model where the dependent variable is logarithm of unit sales and the independent price variables are logarithm of prices.

I start by estimating each functional form for each brand using OLS and test for heteroscedasticity, multicollinearity, and serial correlation and correct for them if detected (see Sethuraman, 1995, for details). From the cross-price coefficients estimated from these models, I compute the cross-price elasticities. In the linear model, elasticity is computed by multiplying the coefficient with mean actual price and dividing by mean sales. In the semilog model, elasticity is computed by multiplying the coefficient with mean price. In the double-log model the price coefficient is itself an estimate of the price elasticity.

1.3. Selection of cross-elasticities

In all, there are seventy-one brands whose sales are estimated using three functional forms resulting in 213 sales equations. The correlation between actual sales and predicted sales (a measure of model fit) for each of the 213 brand sales equations is given in an appendix in Sethuraman (1995). Of these, twelve are excluded because the model fit is inferior to alternate functional forms for that brand (a model is considered inferior in fit if the correlation between actual and predicted sales from the model is below the correlation from an alternate model by more than .05). In the other cases, all functional forms provide more or less the same fit and their estimates are retained. The correlations between actual and predicted sales for these 201 models range from .5 to .99 with an average of .82.

Of the seventy-one brands, fifty-five are national brands (NB) and sixteen are private labels (PL). Our focus is on the effect of price cut of each of the fifty-five national brands on the sales of private label in their market (we denote this effect as $NB \rightarrow PL$) and the effect of the private label price cut on each of the fifty-five national brands in the market ($PL \rightarrow NB$). Thus there are fifty-five brand-pairs in which the effect of national brand price cut on private label sales can be measured, and fifty-five brand-pairs in which the effect of private label price cut on national brand sales can be measured. For each brand-pair, the cross-price elasticities are estimated using three functional forms.

An important criterion for a good estimate is that it should have the correct sign. Conventional economic theory suggests that a brand's price cut would decrease a competing brand's sales—that is, the cross-price elasticity would be nonnegative. Hence, I retain all elasticities that are nonnegative. In 100 of the 110 brand-pairs (55 $NB \rightarrow PL$ + 55 $PL \rightarrow NB$) there is at least one cross-price elasticity estimate that has the correct (nonnegative) sign. In nine of the remaining ten brand-pairs in which the cross-price elasticities are negative, the magnitudes of the elasticity are small (less than .7) and statistically nonsignificant ($t < 1.2$). So I set the elasticity estimate to zero. This approach is somewhat consistent with the study of Allenby (1989) that argues for constrained estimation to obtain correct signs. A similar approach has been used in earlier meta-analysis studies (Sethuraman and Tellis, 1991). One brand-pair with significant negative cross-elasticities is deleted.

Using this procedure, I obtain 137 $NB \rightarrow PL$ cross-price elasticity estimates from fifty-five brand-pairs for analyzing the effect of national brand price cut on private label sales and 124 $PL \rightarrow NB$ estimates from fifty-four brand-pairs for analyzing the effect of private label price cut on national brand sales. In order to maximize the amount of available information used, following Farley and Lehmann (1986), I use all the estimates in our analysis and consider each estimate an observation.

The method of treating multiple elasticity estimates for the same brand-pair as separate observations could lead to problems of duplication and lack of independence. To account for duplication, I weight the observations by the number of replications. That is, where there are valid cross-elasticity estimates for a brand-pair from two functional forms, I weight each of those estimates by 0.5, and where there are three valid elasticity estimates for a brand-pair I weight them by .33. If there is only one valid estimate, I weight them by 1. As for lack of independence arising from analysis of multiple observations from a study, Farley and Lehmann (1986, p. 106) and Hunter and Schmidt (1990, p. 452) point out that the problem may not be very serious if the number of replications relative to the total number of observations is small. In our case, the 261 elasticities come from 109 brand-pairs (average of 2.3 estimates per brand-pair).

2. Meta-analysis of cross-price elasticities

2.1. Overall analysis

The cross-price elasticities range from 0 to 2.12. The weighted average of the 261 cross-price elasticities is .54 (s.d. = .30). The average cross-price elasticities for each product-chain are provided in Table 1. The average cross-elasticities vary by product and store. Tuna and orange juice appear to have large cross-price elasticities, and the percentage sold on deal for these products is also relatively high. The average own-price elasticity is 3.23 across all brands, 3.17 for national brands, and 3.43 for private labels.

For understanding whether, overall, national brand price reductions hurt private-label sales, I analyze the cross-price elasticities from the private label sales models. $\eta(NB \rightarrow PL)$. They measure the percent change in private label sales for 1 percent change in national brand price. There are 137 cross-price elasticities from sixteen product-chains. The cross-price elasticities range from 0 to 1.9. The distribution of cross-price elasticities is given in Figure 1. A majority of the price elasticities are in the range of .2 to .8. The weighted average cross-price elasticity is 0.56 (s.d. = 27) indicating that, across these sixteen product-chains and fifty-five brands, a 1 percent price reduction by national brands results in an average .56 percent decrease in private-label sales. The mean elasticity estimate is significantly greater than zero ($t_{54} = 2.07, p < .05$). Twenty-nine (53 percent) of the fifty-five national brands have significant cross-price elasticities in at least one functional form as evidenced by a t-value greater than 1.67 (critical t for one-tailed test at the 95 percent confidence level). These findings indicate that price reductions by about half the national brands do significantly influence private-label sales and that the effect cumulated across all brands is significantly greater than zero.

For understanding whether private-label price reductions hurt national brand sales, I analyze the cross-price elasticities from the national brand sales models, $\eta(PL \rightarrow NB)$. They

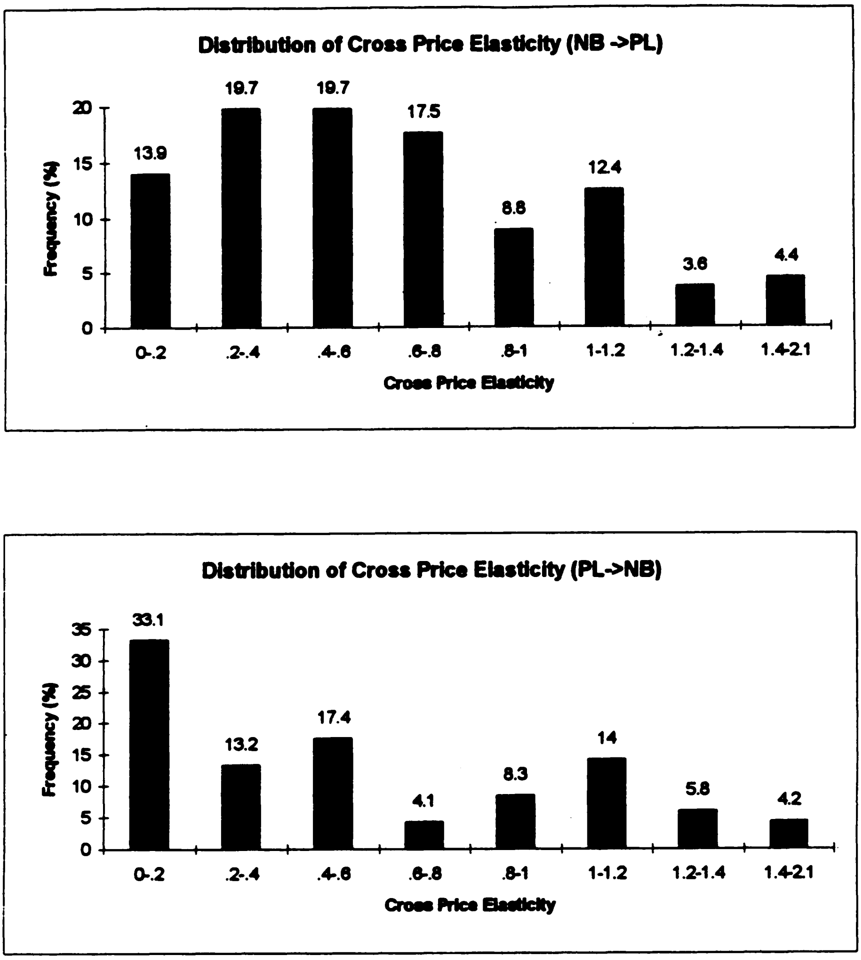

Figure 1. Distribution of cross-price elasticity.

measure the percent change in national brand sales for 1 percent change in private label price. There are 124 cross-price elasticities from sixteen product-chains. The cross-price elasticities range from 0 to 2.12. The distribution of price elasticities is given in Figure 1. Compared to the distribution of $\eta(NB \rightarrow PL)$, there is greater variation in these cross-price elasticities. In particular, in about one-third of the cases, the elasticity is close to zero $(0 - .2)$ and in about 14 percent of the cases, the elasticities are as high as 1. The weighted average cross-price elasticity is 0.51 (s.d. = .33) and is significantly greater than zero only at the 10 percent level ($t_{53} = 1.55, p < .10$). Thirteen (24 percent) of the fifty-four brand-pairs have significant cross-price elasticities ($t > 1.67$) in at least one of the models. The difference in mean cross-price elasticities of $NB \rightarrow PL$ and $PL \rightarrow NB$ (.05) is not significantly greater than zero.

These findings indicate the following. Based on mean cross-price elasticity across all brands, we cannot state that the effect of national brand price cut on private-label sales is greater than the effect of private-label price cut on national brand sales. However, the elasticity estimates relating to effect of private-brand discounts are less stable, which leads to fewer cases where the cross-price effect is statistically significant. Furthermore, there is greater variation in the mean effect of private-label price cuts: private-label price cuts have little impact on several national brands and a large impact on others. What type of national brands hurt private-label sales more, and what type of national brands are hurt by private label price cuts? The next section addresses this question.

2.2. *Brand characteristics that influence cross-elasticities*

We investigate two national brand characteristics—brand market share and relative price. National brands with large market shares are generally the more popular brands and spend heavily on advertising. The awareness level, and hence the number of consumers who are likely to include the brand in their consideration set, is likely to be greater for these brands (Hauser and Wernerfelt, 1989). Because of their popularity, they would also have greater drawing power. Hence, when these brands discount, more private-label consumers are likely to switch, resulting in larger cross-price elasticity $\eta(NB \rightarrow PL)$. On the other hand, private-label discounts are less likely to affect the national brands with larger market share. Brands with large market share possess high market power because of their popularity and generally large advertising levels (Porter, 1976; Bolton, 1989). They are therefore insulated from incursions due to private-label discounts. The small market-share brands have little market power and are more vulnerable to store brand manipulations (Stern, 1966). Hence, we expect $\eta(PL \rightarrow NB)$ to be smaller for large-share national brands.

The influence of brand relative price is not as clear. On the one hand, brands with higher relative price are likely to be perceived to be higher in quality. When these brands promote, private-label consumers are likely to perceive greater value and switch resulting in larger cross-price elasticity $\eta(NB \rightarrow PL)$. On the other hand, the national brands with lower relative price (closer in price to those of private labels) are more likely to be in the consideration set of private-label consumers, hence price reductions by brands with lower relative price are likely to cause a greater reduction in private-label sales.

However, it is reasonable to expect that the consumers of the higher-priced national brands would be less likely to switch to the store brand even when the store brand is promoted because they would perceive a large quality difference. Hence, $\eta(PL \rightarrow NB)$ would be smaller for higher-priced brands.

National brand market share is computed for each national brand in each product and store by dividing the total unit sales of the brand across 104 weeks by the total sales of all brands in the category. Relative price for each national brand is computed by dividing the mean regular price of the brand by the mean regular price of the private label. The average national brand market share is 21.3 percent, the average relative price is 1.31, and the correlation between national brand share and relative price is $-.28$.

To test if the cross-price elasticity (η) varies systematically with market share and relative price, after accounting for variations due to other factors such as product and chain differences, I estimate a multiple regression model. The independent variables are the two brand characteristics, and indicators or dummy variables representing product, chain, and functional form differences. There are five product dummies representing six products, two chain dummies for the three chains, and two functional form dummies for the three functional forms.

The results of the OLS regression are reported in Table 2.[1] The R^2 for the model is .24 (adj. $R^2 = .18$, $F = 3.61$, $p < .05$). The coefficient of national brand share is positive and statistically significant as expected, and the coefficient of national brand relative price is negative and significant.

I estimate a similar model to understand the brand characteristics that influence the effect of private-label price cut on national brand sales. In this model, the independent variables are the same as in the previous model, but the dependent variable is $\eta(PL \rightarrow NB)$ instead of $\eta(NB \rightarrow PL)$. The results of this regression are also reported in Table 2. The R^2 for the model is .39 (adj. $R^2 = .33$, $F = 6.56$, $p < .01$). The coefficient of national brand share and the coefficient of relative price are both negative and significant, as hypothesized.

These findings indicate that national brands with large market share have significant influence on private-label sales but they are less likely to be affected by private-label price cuts. National brands with lower relative price are both more likely to influence private-label sales and more likely to be influenced by private-label price cuts.

The regression results also indicate systematic variations in cross promotional price elasticities due to product and store differences but not due to the functional form used for estimation. Compared to the cross-price elasticity for flour, which is generally considered a commodity product with little perceived quality differential, bathroom tissue and fabric softener have significantly lower cross-price elasticities (negative signs) in both models.

Table 2. Regression results.

Dependent Variable	$\eta(NB \rightarrow PL)$		$\eta(PL \rightarrow NB)$		$\eta(NB \rightarrow PL) - \eta(PL \rightarrow NB)$	
Independent	Est.	*t*-Stat.	Est.	*t*-Stat.	Est.	*t*-Stat.
Intercept	1.27	2.79[a]	2.89	5.66[a]	−.53	−.91
NB share	.006	1.92[a]	−.023	−6.46[a]	.03	5.20[a]
NB relative price	−.47	−1.78[b]	−1.03	−3.41[a]	−.47	−.54
Bath tissue	−.39	−2.58[a]	−.31	−1.81[b]	−.30	−1.18
Fabric softener	−.28	−1.78[b]	−.23	−1.28	−.20	−.69
Margarine	−.12	−.92	−.33	−1.20	.25	1.15
Orange juice	.11	.64	−.09	−.47	.04	.15
Tuna	.05	.33	−.07	−.42	−.02	−.07
Flour (base)	—	—	—	—	—	—
Chain A	−.02	−.18	−.48	−4.75[a]	.50	3.07
Chain B	−.01	−.06	−.37	−3.6[a]	.53	3.24
Chain (base)	—	—	—	—	—	—
Linear model	−.10	−1.2	−.05	−.56	−.01	−.09
Semilog model	−.06	−.79	−.05	−.49	−.04	−.24
Double log (base)	—	—	—	—	—	—
R^2 (adj. R^2)	.24 (.18)	$F = 3.66^a$	.39 (.33)	$F = 6.56^a$	.41 (.34)	$F = 5.63^a$

Notes: *NB* = national brand, *PL* = private label.
a. $p < .05$.
b. $p < .10$.

2.3. *Matched-pair analysis*

In the previous sections, I have separately analyzed the 137 $\eta(NB \rightarrow PL)$ observations and the 124 $\eta(PL \rightarrow NB)$ observations. In this section, I analyze the two effects together by matching an $\eta(NB \rightarrow PL)$ observation with the $\eta(PL \rightarrow NB)$ observation that comes from the same national brand and private-label and estimated using the same functional form.[2] There are 102 matched pairs.

The mean difference in cross-price elasticities $[\eta(NB \rightarrow PL) - \eta(PL \rightarrow NB)]$ from these matched pairs is .06. The correlation between $\eta(NB \rightarrow PL)$ and $\eta(PL \rightarrow NB)$ is $-.17$ ($p < .10$). The negative correlation suggests that national brands that affect private-label sales with their price cuts are less likely to be affected by private-label price discounts. To investigate for what type of brands the difference in cross-price effects is high, I run a regression model with $\eta(NB \rightarrow PL) - \eta(PL \rightarrow NB)$ as the dependent variable. The independent variables are the same as the ones used in the separate analyses.

The regression results are reported in Table 2. The R^2 for the model is .41 (adj. $R^2 = .34$, $F = 5.63$, $p < .01$). The coefficient of national brand share is positive and significant while the coefficient of relative price is not. These results are consistent with results from the separate analyses. In the separate analyses, national brands with large market share have greater effect on private-label sales but are affected less by private-label price cuts—that is, greater $\eta(NB \rightarrow PL)$ and smaller $\eta(PL \rightarrow NB)$. Consistent with this, in the matched-pair analysis, we find the coefficient of national brand share to be positive and significant indicating that the difference in elasticties is greater for large-share national brands. On the other hand, national brands with lower relative price have greater effect on private-label sales and are affected more by private-label discounts—that is, larger $\eta(NB \rightarrow PL)$ and larger $\eta(PL \rightarrow NB)$. Hence the coefficient of national brand price in the difference model is not significant.

3. Conclusion

3.1. *Summary and discussion of results*

Overall, national brand price cuts do significantly influence private-label sales. The average cross-price elasticity is .56 and about 50 percent of the national brands significantly reduce private-label sales. There is mixed evidence for the effect of private-label price cuts on national brand sales. The mean cross-price elasticity of private-label price cut is .51—not significantly different from .56. Hence, based on aggregate cross-price elasticity across all brands, we cannot state that the effect of national brand price cut on private-label sales is greater than the effect of private-label price cut on national brand sales. However, the elasticity estimates for the private brands are less stable, which leads to fewer cases where the cross-price effect is statistically significant. One possible reason for this may be that there are multiple national brands and one store brand in a category, so the effect of private-label price cuts are spread over several national brands resulting in lower stability. Furthermore, there is greater variation in the mean effect of private-label price cuts: private-label price cuts have little impact on several national brands and a large impact on others. Additional analysis reveal some brand characteristics that influence cross-price elasticities.

National brands with large market shares are likely to take more sales away from private labels by discounting than small share national brands, but they are less likely to be affected by private-label price cuts. These findings support the market share = market power theory. Large market-share brands possess market power because of their high awareness and advertising levels and popularity. Hence, when they promote, private-label consumers buy these brands. But because of their market power, they are insulated from private-label price cuts.

These findings are particularly relevant in today's market as manufacturers of leading brands (such as Kellogg's and Procter & Gamble) attempt to stem private-label growth. In our data, the average cross-price elasticity of national brand price cut on private-label sales for the sixteen leading national brands (national brands with the greatest market share in a category is .67. The average cross-price elasticity of private-label price cut on sales of these leading national brands is .32, with the elasticities less than .2 in about 50 percent of the cases. Hence, leading national brands can use price cuts to draw sales from private labels, but they are less affected by private-label discounts.

However, the lack of private-label price effects on sales of large-share brands cannot be interpreted to imply that these national brands do not have to worry about private labels. It simply suggests that price changes by private labels do not affect national brand sales *in the short term*. So national brand managers need not be concerned about private-label discounts. They still need to be concerned about the presence of private labels and their ability to penetrate the market.

National brands with lower relative price influence private-label sales more and, in turn, are affected more by private-label price cuts. In all our data sets, private label is the lower-priced brand. Hence, national brands with lower relative price are the ones that are priced close to the private labels. Our finding indicates that price competition is the greatest between the private label and a national brand that is closer in price to the private label. One possible reason is that brands that are closer in price to those of private labels are more likely to be in the consideration set of private-label consumers and hence compete more on the basis of price. Premium national brands compete less on the basis of price with private labels because they target a separate segment.

3.2. *Limitations and future research*

I have analyzed 261 cross-price elasticities from sixteen product-chain combinations. I believe the results would hold for other frequently purchased grocery products and other supermarket chains. I also recognize that there may be other brand or product characteristics besides the ones examined in this paper that may influence cross-price effects. Future research can analyze more product categories in more retail chains and examine additional brand or product characteristics. Future research can also examine the types or segments of consumers who are likely to switch between national brands and private labels (Lemon and Winer, 1993), especially using scanner panel data. Our store-level analysis is based on aggregate sales and may differ from household-level results.

One possible reason that several national brands do not decrease private label sales with their price cuts may be because they do not reduce the price low enough for the private-label consumers to switch. For instance, some private-label consumers (price-shoppers) may switch only if the discounted price of the national brand equals the price of private label. Future research can investigate the possible existence of such price "threshold" effects.

Finally, I analyze only short-term changes in sales due to short-term price cuts. Hence, the results have to be interpreted accordingly. How price changes affect sales in the long term is an interesting and useful area for future research.

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Notes

1. Diagnostic tests do not indicate problems of heteroscedasticity ($\chi^2_{55} = 56.1, p > .4$) or multicollinearity (variance inflation factors less than 5, condition indices less than 10). Because all brand sale equations in a category have the same set of independent variables, use of SUR (Seemingly Unrelated Regression) is not necessary.
2. I thank Don Lehmann for suggesting this analysis.

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