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In a paper in the March 2004 *American Economic Review*, Justine Hastings studies the acquisition of a sizable independent gasoline retailer, Thrifty Oil Company (Thrifty), by a vertically integrated refiner/retailer, ARCO. She employs a difference-in-differences approach on a panel of station-specific prices to examine the price effects at competing stations of this transaction. She finds that the loss of an independent marketer increased retail gasoline prices by five cents per gallon, but changes in horizontal concentration and differences in the degree of vertical control ARCO exerted over its newly acquired branded outlets did not affect prices. Further empirical results lend support to a particular model of consumer brand loyalty as the underlying mechanism for the post-acquisition price increase.

These results have several implications. Previous research had generally shown that greater degrees of vertical integration are associated with lower retail gasoline prices.[1] Hastings' results suggest that regulations aimed at restricting refiners' vertical control may in fact be benign. Also, merger enforcement may be an appropriate instrument to protect against price effects from the decline in independents' market share from their acquisition by vertically-integrated refiners. In addition, the merger policy implications may transcend gasoline retailing since brand loyalty drives consumption decisions in a variety of product markets.

The sheer size of the estimated price effect – five cents per gallon amounts to a 50 percent increase in retail margins – along with a desire to better understand the novel mechanism behind it, motivated us to revisit Hastings' analysis.[2] Being unable to acquire her data, we used an alternative source, which in aggregate is very similar to the original data set. While there are differences between the two data sets, the five-cent effect is large enough that we would expect to find an effect of a similar order of magnitude. Ultimately, however, we find an effect of approximately three-tenths of a cent per gallon, which is of little economic (and often statistical) significance. This finding is robust to using various sub-

[1] For examples of papers that find vertical integration pro-competitive or bans on vertical integration being anticompetitive *see*

samples, analysis of higher-frequency data, unavailable in Hastings' data, and whether or not we use clustered standard errors at the station-level.

In addition, we examine the theory of brand loyalty as outlined by Hastings. She employs a modified Hotelling model in which both firms will raise prices after an independent converts to a branded station. We point out that the corresponding increase in gross consumer utility in the model implies that total welfare must also increase as a consequence of rebranding. Furthermore, we find parameter values such that total consumer surplus increases after rebranding, *even* in the face of an across-the-board price increase.

Thus, our empirical results cast doubt on whether ARCO's acquisition of Thriftyour andederue acquisitioniicrease

and December in 1997" (Hastings, p. 321). In other words, there are four data points for each station, two before and two after the transaction.

Being unable to acquire Hastings' data, we examine a panel of station-specific retail prices for gasoline outlets in Los Angeles and San Diego from February 1996 through December 1998 from the Oil Price Information Service (OPIS).[7] Most of our results rely on the 1997 data. OPIS collects the data from fleet card transactions.[8] We use the average weekly price charged at a given outlet for a gallon of regular unleaded gasoline.[9] We also obtained a 1997 census of gasoline station locations from the California Energy Commission (CEC).

It is important to gauge the extent to which the OPIS data differ from the W-L data. Table 1 compares the distribution of brands in the OPIS and W-L data sets against the CEC census, for both Los Angeles and San Diego. Both the OPIS and W-L data under-sample minor brands and independents, as measured by the CEC census.[10] However, the overall distribution of brands in the W-L data better aligns with their actual brand shares as measured by the CEC census. Although the OPIS data set contains more stations, it omits some major brands (specifically, ARCO, Chevron, and Unocal) as well as a number of minor brands and independents.[11] Neither the OPIS nor W-L data set captures price information at rebranded Thrifty stations either before or after the transaction.

To summarize, both the W-L and OPIS data sets provide retail price information at the station-specific level. While the W-L data set is a more representative sample of the true distribution of brands, the OPIS data are available at a greater frequency (*i.e.*, daily or weekly), for a larger number of stations

[7] Professor Hastings denied requests for access to her data set since it is copyrighted. We were not able to purchase the data since the Whitney Leigh Corporation is now defunct.

and for a longer time period; possibly allowing for a more precise estimate of the extent to which the ARCO/Thrifty transaction affected retail gasoline prices.[12]

The ultimate question is to what extent the distribution of brands in the OPIS data could bias our results. We can make a conjecture based on Hastings' Table 3, which documents how the ARCO/Thrifty transaction differentially affected "high-share", "mid-share", and "low-share" brands.[13] (The effect on ARCO's pre-merger stations is estimated separately.) Column (7) of our Table 1 displays Hastings' classification for each brand. Hastings posits that market share is positively correlated with the degree of brand loyalty. Consequently, independents (*e.g.*, Thrifty) compete most closely with low-share brands (*e.g.*, Citgo), so that the ARCO/Thrifty transaction should increase prices the most at low-share branded stations. Indeed, using the W-L data Hastings finds that prices increased at low-share, mid-share, and high-share brands by approximately \$0.07, \$0.05, and \$0.03 per gallon, respectively. Consistent with its low-price strategy, the estimated effect on ARCO is similar to the low-share group.

Based on the OPIS data over-sampling high- and mid-share brands at the expense of low-share brands and ARCO, we expect our estimate of the transition's effect on all stations to be lower than that estimated by Hastings. However, this difference should be small, since the effect on high- and mid-share brands was \$0.03 and \$0.05. Additionally, the OPIS data over-samples a low-share brand, Citgo, in San Diego, especially relative to the W-L data. So the effect of the distribution of brands in the San Diego OPIS data is ambiguous. For the OPIS *pooled* sample (consisting of both Los Angeles and San Diego), we expect that the difference between the OPIS and W-L data sets may still lead to a slightly smaller

Following Hastings, we begin our replication attempt by considering simple before-and-after changes in the average price of gasoline, measures across all stations in the OPIS samples, around the time of the ARCO/Thrifty transaction. Figure 1 presents three graphs corresponding to the pooled, Los Angeles, and San Diego OPIS price data. The graphs reflect the average weekly retail price in each area as measured during the last week of February, June, October, and December of 1997.

Despite the use of differing price series, it is striking how closely the graphs presented in Figure 1 match the general shape and levels of the corresponding graphs, Figure 1, presented in Hastings' study. In each city the average gasoline price level for each set of stations peaks in October 1997, with the rise and fall more pronounced in Los Angeles than in San Diego. Hastings' graphs exhibit the same features. However, the June price levels in our OPIS data exceed those in the W-L data by a few cents per gallon, within each city and for both sets of firms.

When considering the trends in the time series of those stations that competed against a Thrifty versus those that did not in the OPIS data, there is substantially less agreement with the W-L data. For Los Angeles the two series track each other very closely, with the prices charged by Thrifty-competing stations *always* (*i.e.*, both before *and* by \$0.04-\$0.06 per gallon. By contrast, the San Diego OPIS data show that Thrifty competitors were only \$0.01 below and then \$0.01 above other stations, suggesting an effect of only \$0.02 per gallon.

The graph for the pooled sample for the OPIS data presented in Figure 1 again shows no effect on the relative positioning of the control or treatment group price series over the period. This result is not surprising given that most of the observations in the pooled sample come from Los Angeles.

C. Econometric Analysis

In this section we consider several empirical specifications based upon the econometric research design employed by Hastings. Specifically, we adopt Hastings' fixed-effects or difference-in-differences approach to identifying the impact of Thrifty conversions on market prices, both in the aggregate and by individual brand, and for both the pooled and individual OPIS city samples.

Table 2 presents the results of estimating the following regression:

$$p_{i,t} \quad N_i \quad C\,\mathrm{Conversion}_{i,t} \quad \sum_{j\ 1}^{N}\sum_{k\ 1}^{T\ 1} {}_{jk}\, I^{j}_{i}\, F^{k}_{t} \quad H_{i\,t} \quad U \quad (1) \quad F$$

The above model is identical to Hastings' except in two regards.[14] First, in classifying Thrifty competitors she uses actual driving distance of one mile, whereas we employ the simpler one-mile radius "as the crow flies." Our methodology likely counts more stations as Thrifty competitors than Hastings' would. If the transaction's effect diminishes with driving distance, our results would be biased downward. Since Hastings reported, in footnote 15, that half mile changes in the definition of which stations competed with a Thrifty did not significantly change the results, this small difference in should not explain the gap between our estimates and Hastings'. Second, Hastings includes a dummy variable indicating whether ARCO operated the station after the rebranding. We lack this information. But Hastings finds that the point estimate of the coefficient on this variable is less than one cent per gallon and is statistically insignificant.

Table 2 presents our results separately for the pooled, Los Angeles, and San Diego OPIS samples as well as Hastings' results for the pooled sample. The estimated standard errors used in constructing the reported *t*-statistics are obtained using the Huber/White heteroskedasticity-consistent covariance matrix estimator. The models appear to fit the data relatively well, with the *R*-squared ranging from 0.90 to 0.92. The coefficient estimates for the individual city-month interactions are economically relevant, ranging from $0.02 to $0.13 per gallon. These estimates closely resemble Hastings' findings using the W-L data, which is not surprising given the similarity of the time trends in the graphical analyses.

By contrast, the coefficient estimates pertaining to the conversion variable in our Table 2 differ substantially from Hastings'. Hastings' regression results for the pooled W-L sample indicate that the loss of Thrifty as a competitor is associated with a price increase of $0.05 per gallon. For the OPIS pooled sample, the coefficient estimate on the conversion variable is likewise negative – however, it is an order of magnitude smaller at just two-tenths of a cent per gallon and not significantly different from zero at conventional levels of statistical significance. For the Los Angeles OPIS sub-sample, the point estimate

[14] A third, technical, difference is that Hastings considers whether each station competes with *any* independent, not just Thrifty. However, she notes in her footnote 18 (p. 324) that the only source of variation in this variable comes from the Thrifty conversion. Thus, the station-specific fixed effect should capture any effect of competition from non-Thrifty independents, with no effect on our results.

of the conversion indicator is actually *positive* – indicating a price *decrease* after the acquisition – but quite small, less than one-tenth of a cent per gallon. For the San Diego OPIS sub-sample, the point estimate is found to be negative and statistically significant at the 1 percent level. The magnitude of the estimated conversion effect is $0.01 per gallon. This estimate comports with our graphical analysis Figure 1(c) but is smaller than the effect in Hastings' San Diego graph, $0.04 to $0.06 per gallon. This estimate is also less than our expected effect given that the San Diego OPIS data over-sample a low-share brand, Citgo.

As mentioned earlier, the OPIS data possess the benefit of being available at a higher frequency and also allows for the effects of the ARCO/Thrifty transaction to be considered over a longer time period than Hastings' W-L data. Table 3 presents the results of estimating the equation (1) where t spans all individual weeks. This "two-way" fixed-effects specification controls for both station-level (group) and week (period) fixed-effects within each sample. The week fixed effects control for any week-specific unobserved factors that influence retail gasoline prices symmetrically across all stations in a city. This equation was estimated separately for Los Angeles, San Diego and the pooled samples. In Table 3, the results strongly suggest that it was much smaller than that suggested by Hastings' study using the W-L data. The bottom row of Table 3 shows that statistical significance diminishes considerably when we cluster standard errors to control for within-station autocorrelation. The conversion effect's statistical significance tends to diminish as well when estimating the model using all three years of data.[15]

As discussed above, Hastings concludes her analysis by disaggregating the effect of the transaction by high-, mid-, and low-share brands to see if the pricing patterns support her theory of product differentiation with brand loyalty. While these results are generally consistent with the theory – the magnitude of the effect increases as we move from high- to mid- and then low-share brands – the evidence suggesting significant *across* group differences is rather weak. The coefficient estimate on the mid-share variable is not statistically different from that on the low-share variable, while the coefficient estimate on the high-share variable is statistically different from that on the mid-share variable at the 10 percent level. It seems that these results would have been even weaker had Mobil been placed in the high brand group, which would have been more consistent with its market share as reported in footnote 25 of Hastings (2004).

We also tested this theory with the OPIS data. Table 4 presents the results of estimating equation (1) by individual brand (*i.e.*, where the dependent variable is the average weekly price charged by each station) for each of the brands for which sufficient price observations are available. Like the previous regressions, these models compare the prices in a control group, the branded stations not near a rebranded Thrifty, to a treatment group. These regressions use data for all weeks in 1997 controlling for both station and week fixed effects. We present *t*-statistics using Huber/White robust standard errors as well as station-level clustering. In Table 4, we order the brands according to Hastings' classification as high- (Shell), mid- (Mobil and Texaco), or low-share (Citgo) brands.

[15] Additional results showing the robustness with respect to the treatment of stations with mis-classified premium and mid-grade gasoline prices are available, currently, Table A-1. The results did not change by more than one-tenth of a cent per gallon when we either dropped those stations or when we used the minimum daily price instead of the average of the daily prices when aggregating the daily prices to weekly prices.

The rightmost column of Table 4 affords the most direct comparison with Hastings' Table 3. That the level of the estimated effect is much smaller in our table is unsurprising. However, the pattern across brands apparent in Hastings' table is not present in ours. Based on her results using the W-L data, and consistent with her theory of brand loyalty, the magnitude of the coefficient estimates should increase as we move down the table from a high-share to mid- and then low-share brands. However, this pattern does not emerge in our table, for the pooled sample or either of the city sub-samples. In fact, the estimate pertaining to the low-share brand, Citgo, is positive at $0.01 per gallon. This latter result is directly opposite to that predicted by Hastings' theory as it suggests that the ARCO/Thrifty transaction lowered prices by $0.01 per gallon at a competitor that was "close" to Thrifty in product space. Hastings' theoretical hypothesis and empirical results using the W-L data suggest that this type of firm should experience the largest price increase upon the re-branding of an independent competitor.

III. Theoretical Model of Brand Preferences

Hastings posits that a demand structure consistent with her empirical results using the W-L data involves heterogeneous preferences over brands of gasoline. In this section we employ such a model to demonstrate that even when rebranding leads to price increases, its effect on welfare remains ambiguous.

Consider a Hotelling model of product differentiation, with two firms *A* and *B* at the endpoints of a line with length ".[16] Three consumer types are uniformly distributed along the line with unit demands. Proportions D and E are brand-loyal to *A* and *B* respectively, while a proportion J views gasoline as an homogenous product. With transport costs of *t* per unit, if both firms find it profitable to sell to brand-loyal as well as non-brand-loyal customers then equilibrium prices will be

$$p_A \quad c \quad t \quad \frac{2t}{3} \quad B \quad , p_B \quad C\,c \quad "\,t \quad \frac{2t}{3H} \quad B2 \quad C$$

[16] In an earlier working paper, Hastings (2002, p. 30) specifies precisely this model.

both before and after rebranding. Consequently, the post-rebranding price increases are small. So long as the reservation price *r* is high – but not so high to induce firms to ignore the non-brand loyal segment – the gain to *E*-type consumers outweighs the effect of the price increase on *D*- and *J*-type consumers.

Thus, while Hastings' theoretical model does show that rebranding leads to across-the-board price increases, this result arises only because some consumers receive higher gross utility from purchasing branded product. Conse

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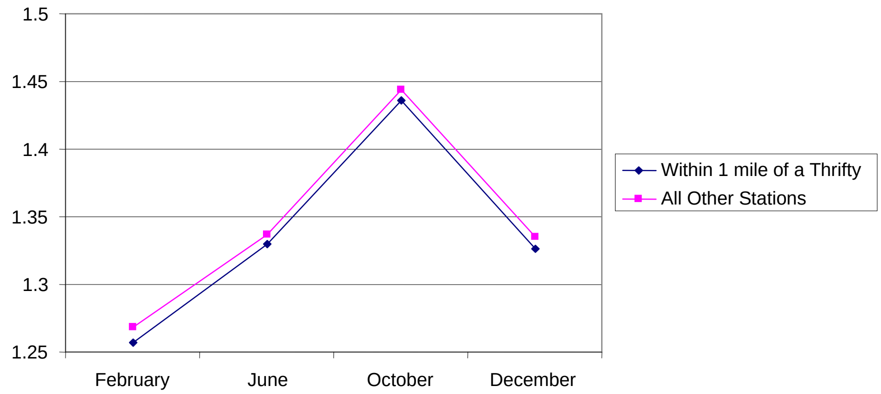
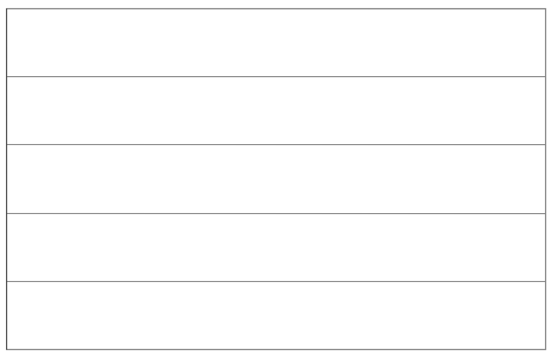
FGURE 1 TREATMENT AND CO GRPANRO PO SIES, LO ANGELES, AND SAN DEGO SAMPLES



(a) PO SAMPLE LED



(b) LO ANGELES SAMPLE

(c) SAN DEGO SAMPLE

Notes: Data are derived from

m onth

Table 1

Comparison of Brand Shares Across Alternative Samples, 1997

Brand	(1) CA Energy Comm. (LA)	(2) CA Energy Comm. (SD)	(3) Whitney Leigh (LA)	(4) Whitney Leigh (SD)	(5) OPIS (LA)	(6) OPIS (SD)	(7) Hastings (2004) Brand Type
Alliance	0.04%						
American Gas	0.35%						
ARCO	11.41%	11.16%	19.41%	13.21%			
Chevron	12.23%	10.34%	17.84%	17.61%	0.35%	0.48%	High
Circle K	0.13%	3.13%					Low
Citgo	0.30%	8.70%			0.32%	18.69%	Low
Exxon	0.13%					2.42%	Mid
Fastrip	0.04%						
Mobil	13.75%	8.44%	15.88%	13.21%	51.19%	23.60%	Mid
Olympic	0.04%						
Shell	14.05%	12.38%	14.12%	17.61%	37.79%	29.81%	High
Sinclair					0.49%		
Texaco	5.03%	10.34%	8.43%	12.58%	9.88%	24.99%	Mid
Thrifty	4.34%	4.76%					
Ultramar	0.30%	1.50%					Low
Unbranded	19.43%	19.73%					
United Oil	1.65%						
Unocal	15.13%	8.71%	12.55%	11.95%			High
USA	0.78%	0.54%					
World Oil	0.87%	0.27%					
Minor Brands			5.25%	8.18%			
Independents			6.52%	5.66%			
Total Stations	2354	806	510	159	582	209	

Notes: Brand shares reported Columns (3) - (4) are from Hastings (2004) Table 1. Brand shares may not sum to 100% due to rounding.

Table 2
Estimated Effect of Thrifty Station Conversions on the Market Price for Retail Regular-Grade Gasoline, 1997

Variable	Los Angeles Sample	San Diego Sample	Pooled Sample	Hastings (2004) Pooled Sample
Conversion	.000334	-0.011**	-0.002	-0.050**
	(0.16)	(2.99)	(1.39)	(4.95)
LA*February	-0.051**		-0.050**	0.018**
	(31.44)		(31.55)	(2.77)
LA*June	0.022**		0.023**	.0243**
	(14.75)		(15.64)	(3.74)
LA*October	0.131**		0.131**	0.139**
	(95.65)		(95.61)	(21.72)
SD*February		-0.111**	-0.114**	-0.085**
		(40.64)	(43.08)	(23.61)
SD*June		-0.049**	-0.051**	-0.030**
		(20.06)	(22.54)	(8.44)
SD*October		0.048**	0.048**	0.055**
		(22.97)	(22.86)	(15.14)
Constant	1.292**	1.445**	1.333**	1.362**
	(1204.10)	(878.21)	(1473.33)	(47.45)
N	2298	832	3130	2676
R^2	0.91	0.92	0.90	0.72
F-statistic (H_O : All slopes = 0)	5330.49**	1390.09**	3845.68**	

Notes: The dependent variable is the average weekly retail price for regular unleaded gasoline by station for the last week in February, June, October, and December of 1997. Absolute values of robust t-statistics appear in parentheses. All models include full sets of station-specific dummies. The symbols * and ** indicate statistical significance at the 5% and 1% levels, respectively.

Table 3

Estimated Effect of Independent Retail Gasoline Station Up-Branding: All Weeks in 1997 and 1996-1998

	Los Angeles Sample		San Diego Sample		Pooled Sample	
	(1) 1997	(2) 1996-1998	(1) 1997	(2) 1996-1998	(1) 1997	(2) 1996-1998
Conversion	-0.003**	0.0006	-0.006**	-0.010**	-0.004**	-0.002**
	(4.87)	(1.58)	(5.94)	(13.36)	(7.30)	(5.86)
Constant	1.172**	1.185**	1.289**	1.240**	1.19**	1.20**
	(1041.63)	(873.46)	(488.01)	(380.26)	(1139.26)	(973.35)
N	29,840	85,767	10,733	30,685	40,573	116,452
R^2	0.93	0.91	0.89	0.84	0.65	0.87
$F(H_O$: All slopes = 0)	8,260.55**	13,212**	1,522.29**	3,249.84**	4,937.89**	3,569.49**
Absolute value of t-statistic of Conversion with station-level clustering	1.63	0.32	1.62	2.59*	2.28*	1.16

Notes: The dependent variable is the OPIS average weekly retail price (measured in dollars per gallon) for regular unleaded gasoline by station. Absolute values of robust t-statistics appear in parentheses. Estimated coefficients and t-statistics for station dummies and week dummies are not reported. The symbols * and ** indicate statistical significance at the 5% and 1% levels, respectively.

Table A-1

Robustness of Estimated Effect of Independent Retail Gasoline Station Up-Branding   Minimum Priœ and Excluding Some Stations.

	Los Angeles Sample			San Diego Sample			Pooled Sample		
	(1) 1997	(2) Minimum	(3) Excluding	(1) 1997	(2) Minimum	(3) Excluding	(1) 1997	(2) Minimum	(3) Excluding