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ASYMMETRIC DYNAMIC PRICING IN A LOCAL GASOLINE RETAIL MARKET*

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Asymmetric-price adjustment is a common phenomenon in many markets around the world, particularly in retail gasoline markets. This paper studies the existence of this phenomenon in the retail gasoline market in the city of Santiago, Chile, using a data set of weekly gas station prices that covers a period of almost four years. We found that prices adjust asymmetrically, and the asymmetry is different for branded gas stations and unbranded stations. In addition, we found that the asymmetry for high-margin stations is statistically equivalent to that for low-margin stations. This evidence is suggestive of collusion as a rationale for the asymmetric pricing policy observed.

I. INTRODUCTION

HOW GASOLINE PRICES ARE SET has been a controversial issue in developed countries such as the U.S.A., the U.K., Canada, the Netherlands and Germany for a long time. In particular, it has been argued that retail distribution companies wield their market power not only to set prices above marginal costs, but also to increase prices rapidly when faced with cost increases; however, when faced with cost decreases, prices are slowly adjusted, allowing for an even larger markup over a short period of time. This behavior is often referred to as the rockets and feathers phenomenon.

This paper documents the existence of the rockets and feathers phenomenon between gasoline prices at the gas-station level (hereinafter retail prices) and gasoline prices at refinery level (hereinafter refinery prices) in the city of Santiago, Chile. The data set used consists of a time-series panel of 44 gas stations located in Santiago that were observed weekly between the first week of March 2001 and the second week of August 2004.[1] This data set contains information regarding prices, geographical location and brand identity for each gas station.

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[1] These represent about 10% of the total number of gas stations located in Santiago.

There are two main rationales for the rockets and feathers phenomenon for the price relationship studied here: implicit collusion and consumer search theories. The implicit collusion theory as developed by Haltiwanger and Harrington [1991] establishes that prices are more likely to adjust quickly when an expected marginal cost increase lowers collusive profits more than non-collusive profits.[2] Consumer search rationales are provided by Tappata [2007] and Cabral and Fishman [2006].[3] For our purpose here the main feature of these models is that they either require that consumers are uninformed about cost shocks or that firms are uniformed about other firms' cost shocks or both.

The Chilean gasoline market has several unusual characteristics that make the study of the rockets and feathers phenomenon interesting from both a theoretical and empirical point of view. At the theoretical level there are two features that are relevant. First, the refinery price is announced by ENAP (the state-owned oil company) every week on its web site and applies to each wholesale buyer irrespective of identity and volume bought. Second, ENAP announces each Friday the variation that retail prices should face the following Monday if gas stations were to fully transfer cost changes (variations in the refinery price) to the final consumer. This reference price is publicized by all major radio and TV stations and the country's main newspapers. As a result of this, uncertainty among both consumers and competing firms regarding weekly price variations and cost shocks is removed. As a consequence of this, consumer-search rationales are less plausible for the Chilean market in relation to other local gasoline-retail markets, while implicit collusion rationales are made more plausible. The former is due to the fact that search rationales require the existence of either uninformed consumers or uninformed firms or both, while the latter is due to the fact that the weekly public-announcement policy leads to frequent price adjustments (once a week and every week that a cost shock occurs), helps to avoid price wars due to bad luck, and provides a natural focal point on price changes that firms can use to agree on a pricing policy.

At the empirical level there are several features that are important. First, ENAP owns all of the country's refineries and supplies around 85% of total national gasoline consumption. Second, the refinery located in the city of Concón, located 100 miles west of Santiago, is the sole supplier to the Santiago wholesale gasoline market. Third, 97% of oil consumed in Chile is imported and Chilean demand represents a negligible share of world oil demand. In fact, Chile is a textbook example of a price-taker

[2] See, Borenstein *et al.* [1997] and Verlinda [2008] for a detailed discussion of this model.

[3] Lewis [2007] also provides a search rationale with bounded rationale agents. See his paper for a detailed discussion of consumer search rationales.

country.[4] Fourth, gas stations adjust prices once a week – usually on the Monday following the Friday announcement by ENAP. Together with the fact that weekly price data are used, these characteristics minimize the temporal aggregation problem suffered by most studies (see: Geweke [2004]),[5] ensure that the refinery price is a very good proxy for the major cost component in the formation of the retail price for the market studied here than for other local retail markets (such as the one studied by Lewis [2007]). They also minimize the simultaneity in the price determination process.

We found, after accounting for heterogeneity in price adjustment dynamics at the level of gas stations and controlling for cross-section dependence,[6] that there is strong evidence in favor of an asymmetric price adjustment pattern – or rather, retail prices respond faster to a refinery price increase than they do to a refinery price fall. Gas stations on average respond to a one peso refinery price increase per liter by increasing prices by \$1.062 per liter during the first week, while they respond to a one peso refinery price decrease per liter by decreasing prices by \$0.895 per liter. In other words, gas stations increase retail price by 6.2% more than the refinery price increase, while they decrease price by 10.5% less than the refinery price decrease. This initial difference of \$0.167 pesos increases gradually to reach a maximum of \$0.27 six weeks after the initial shock and then declines gradually towards zero until retail prices settle at their estimated long-run response. This asymmetry implies that during the first week a consumer, whose weekly gasoline consumption comes to 40 liters, spends \$26 (U.S. \$0.049 or 6.5%) more in his weekly consumption when the refinery price increases \$10 (U.S. \$0.018) per liter compared with the case in which the rockets phenomenon is absent; whereas he spends \$36.5 (U.S. \$0.079 or 9.1%) more due to the feathers phenomenon when a refinery price decrease of \$10 per liter takes place. When the whole adjustment period is taken into account, the rockets phenomenon results in an extra cost of \$213.9 (U.S. \$0.40) and the feathers phenomenon in an extra cost of \$712.9 (U.S. \$1.35). In other words, the rockets phenomenon results in an extra cost of 4.1% and the feathers phenomenon in an extra cost of 13.7% over the whole adjustment period.[7]

[4] This ensures that simultaneity problems are unlikely to be an important component in the determination of the refinery price.

[5] The temporal aggregation problem refers to the case in which prices vary within a given period of time and the collected data corresponds to a greater period of time. A good example of this problem is Borenstein *et al.* [1997], who, using weekly data, find evidence of asymmetry, while Bachmeier and Griffin [2002] using daily data for the same region and relationship, find no evidence of asymmetry.

[6] These methodological issues have been largely ignored in the rockets and feathers literature.

[7] The fact that the percentage change for the whole period is lower for the refinery price increase is due to the fact that after the fifth week, retail price decreases gradually to the long-run equilibrium.

The pattern found, although asymmetric as in most gasoline markets, is particular with regards to the magnitude of adjustment in the first week following the initial cost shock, as this is quite large in relation to the magnitude found in similar studies. In fact, Lewis [2007] reports that during low-margin periods a one cent increase in the Los Angeles spot market gasoline price leads to a 0.35 cents increase in retail prices in San Diego during the first week, while a one cent decrease results in a 0.1 cent drop in retail prices during the first week.[8] Verlinda [2008] found that a one cent increase or decrease in the Los Angeles spot market gasoline price leads to a 0.65 cents rise or fall in retail prices for South Orange County, California. Regarding the price-response asymmetry, Verlinda and Lewis find that greater asymmetry occurs three weeks after the initial shock. Verlinda finds this to be equal to 0.28 cents, and Lewis finds it to be close to 0.33 cents during high and low-margin periods. This suggests that the link between prices and costs is stronger for the Chilean market than for other local markets studied, and the asymmetry during the first week (0.167 cents) is smaller than the one found in Lewis (0.25 cents), but greater than the one found by Verlinda (0 cents). While it is difficult to explain these differences, it is highly plausible to assume that they are in part due to the uncommon characteristics of the Chilean market regarding consumer awareness of cost shocks and competitors' information concerning cost structures.

With regard to the effect of local market power on the price-response asymmetry documented here, we found that branded stations have a more asymmetric price-adjustment pattern than unbranded gas stations and low-margin stations exhibit a more symmetric adjustment pattern than high-margin stations. However, as in Lewis [2007], this difference between gas stations with different margins is not statistically significant. This evidence suggests that the institutional features of the Chilean market facilitate the adoption of a collusive pricing policy in which ENAP's price change recommendation is a focal point, and collusion is easier to sustain the stronger the brand loyalty.

The rest of the paper is as follows: In the following section a brief review of the empirical literature is presented. Section 2 provides a brief description of the Chilean gasoline market. In Section 3, the empirical methodology is discussed. In section 4, the data is discussed and descriptive statistics are provided. Section 5 reports the results regarding the existence of the rockets and feathers phenomenon. In Section 6, we study the link between local market power and price-adjustment behavior. The last section presents some concluding remarks.

[8] During high margin periods, there is a 0.05 cent increase and a 0 cent drop respectively.

II. THE CHILEAN GASOLINE MARKET

Like most gasoline markets, the Chilean market is composed of four different segments: oil extraction and refining, transportation and storage, wholesale distribution, and retail distribution. Regarding which we shall briefly describe each of these in turn.

ENAP is the only company that refines oil in the country. For the year 2005, ENAP supplied 85% of total national demand, which was just 12.7 million m^3. The Concón refinery, located about 100 miles from Santiago and with a daily production capacity of 16,700m^3, supplies all the wholesale distribution companies located in Santiago.[9]

The gasoline storage system in Santiago is composed of four different storage plants with a total capacity of 270,000m^3. All these plants are located in the same geographical area within Santiago: the Maipú district. The biggest plant is owned and operated by ENAP and has a storage capacity of 171,378m^3. The next biggest is operated by Compap, a company owned by the wholesale distribution companies Copec and Shell, with a capacity of 55,350m^3. The next in size is owned by the wholesale distribution company Esso and has a capacity of 23,336m^3, and the smallest plant is owned by a small local company called JLC and has a capacity of 19,936m^3. All four are connected through a pipeline to ENAP's refinery located in Concón. This pipeline is owned by Sonacol, which is a joint venture between ENAP and the wholesale distribution companies Esso, Copec and Shell.

The wholesale segment is composed of five different companies that mainly buy gasoline from ENAP. Of these five companies, four of them (Copec, Esso, Shell and Repsol-YPF) represent 99% of the wholesale market; the rest is in the hands of JLC. Copec has the largest market share (40%), Shell's market share stands at 26.5% Esso has a share equal to 19.8%, and Repsol-YPF's share is 12.8%.[10]

Lastly, the retail distribution segment consists of approximately 450 gas stations that sell gasoline under the brand of one of the wholesale companies. In particular, for the year 2002, Copec's market share was 51%; Shell had 20% of the market; Esso had 19%; and YPF's share was just 9% of the market. The distribution company JLC has just one gas station located in the eastern part of the city. All gas stations sell non-leaded gasoline at 93, 95 and 97 octane. The average share of each type of gasoline in the total demand for the period considered here was 42%, 30% and 28% respectively.

The relationship between the wholesale and retail segment is not all arm's length. In fact, 7.3% of gas stations are operated by independent

[9] Even though the refining segment is a local monopoly, the possibility of importing gasoline directly from abroad limits its potential market power.

[10] Market share is measured in terms of the number of gas stations and not sales. There are no sales data available.

distributors, while the remaining 92.7% are operated under one of the following contractual schemes: i) 40% are owned by the wholesale distribution companies and operated either by a leessee-dealer or by a franchiser; ii) 16.6% are operated by related companies or subsidiaries of the wholesale companies; iii) 8.5% are independently owned, but operated by one of the wholesale companies through a lease contract or franchise; and iv) 27,6% are leased by the wholesale companies, which simultaneously lease them to independent operators who sell gasoline under the brand of the leasing wholesale company and pay a fixed amount to the wholesale company that supplies the gasoline.

Given that 85% of the gasoline traded in the country is sold by ENAP, its pricing policy is crucial for understanding the price structure in the internal gasoline market. During the period taken into account in this study, from the first week of March, 2001, to the second week of August, 2004, ENAP followed a pricing policy based on a weekly calculated international parity price, which accurately represents the opportunity cost of the gasoline sold at the refinery in Concón. This price corresponds to the price of gasoline in the Gulf of Mexico plus transportation and logistical costs, international trade tariffs, and the importer's mark-up. ENAP sells gasoline in non-discriminatory terms and without a quantity discount to wholesale companies in Santiago, at a price equal to the international parity price plus the excise tax – which is currently equal to U.S. $1 per liter –, a value-added tax of 19%, the adjustments made due to the oil stabilization fund, and the transportation costs corresponding to the use of the pipeline connecting the refinery in Concón with the storage plants located in Maipú. On average, the price paid by wholesale companies located in Santiago can be broken down as follows: 46.5% corresponds to the price paid to the refinery in Concón, 36.8% is excise tax, 5.8% goes to the oil stabilization fund, 0.8% covers transportation costs, and 10.1% is paid in value-added tax.

In short, in each segment market concentration is relatively high compared with most countries and the wholesale and retail segments are closely related. This, together with the fact that wholesale companies face almost identical (if not identical) transportation and storage costs, as well as the same refinery price, implies that the price paid at the Concón refinery is a very good approximation of wholesale costs. Furthermore, the fact that arm's-length trade between segments is rare suggests that the refinery price is also a good proxy of the gas stations' main component of total costs – the wholesale price. It is also clear that the other components of gas stations' total costs differ across stations, yet given the time span of this study and its goal, most of them can be regarded as constant over time. Furthermore, some components vary over time, yet they do not vary with the amount of gasoline sold and as such these components have no effect on marginal costs.

III. METHODOLOGY

The standard econometric model used to test for the presence of the rockets and feathers phenomenon, introduced through the pioneering work of Borenstein *et al.* [1997], is the error-correction model (hereafter ECM). Usually, the price equation is expressed as an autoregressive distributed lag $ADL\,(n, m)$ model, where n refers to the number of lags of the refinery price (W) and m to the number of lags of the retail price (R). The ADL model is extremely flexible since it is capable of capturing complex price dynamics. This is often seen in the following bivariate form:

$$(1)\qquad R_{s,t} = \alpha_s + \delta t + \sum_{i=0}^{n} \beta_i^s W_{t-i} + \sum_{k=1}^{m} \gamma_k^s R_{s,t-k} + \upsilon_{s,t},\ \text{for } s = 1, \ldots, S \text{ and } t = 1, \ldots, T$$

where $R_{s,\ t-k}$ is the retail price set by gas station s in period $t-k$, α_s is the fixed effect, and W_{t-i} is the price at the Concón refinery in period $t-i$.

In its construction, equation (1) allows for short-run price fluctuations that are a function of observed prices and costs in the past. However, economic theory suggests that the relationship between prices and costs should be governed by a long-run relationship in which prices increase with costs. When cost differences follow a random walk, the existence of a long-run relationship results in an econometric problem commonly known as co-integration. In this case, equations at levels such as equation (1) are not capable of identifying the different coefficients on cost variables since W_t differs from W_{t-1} only on the residual of the regression of W_t on $W_{t-\mathrm{i}}$. Furthermore, this creates an omitted variables bias since the residual in (1) is correlated with cost regressors (current and lagged).

Engle and Granger's 1987 representation theorem shows that when the series are co-integrated, this difficulty can be solved by means of re-parameterizing the model in equation (1) using the following identities: $\Delta R_{s,\ t} \equiv R_{s,\ t} - R_{s,\ t-1}$ and $\Delta W_t \equiv \Delta W_t - W_{t-1}$. This transforms the model in levels to a model in first differences by means of substituting $R_{s,\ t}$ for $R_{s,\ t-1} + \Delta R_{s,\ t}$ and W_t for $W_{t-1} + \Delta W_t$.[11] In addition, the model in equation (1) assumes that the response of gas stations to a change in the refinery price is the same whether this increases or decreases. However, the evidence suggests that gas stations usually respond more to cost increases than to cost decreases. Thus, to study the presence of this behavior, known sometimes as short-run asymmetry or amount asymmetry, the model in equation (1) is

[11] For example, for $m=2$ and $n=1$, $\tilde{\beta}_0^s = \beta_0^s$, $\tilde{\alpha}_s = \alpha_s/(\gamma_1^s - 1)$, $\tilde{\delta} = \delta/(\gamma_1^s - 1)$, $\theta_s = (\beta_0^s + \beta_1^s)/(\gamma_1^s - 1)$, and $\lambda_s = \gamma_1^s - 1$.

further generalized by differentiating between positive and negative price changes.[12] This leads to the following ECM:[13]

$$
\begin{aligned}
\Delta R_{s,t} = {} & \sum\nolimits_{i=0}^{I} \tilde{\beta}_i^{s+} \Delta W_{t-i}^{+} + \sum\nolimits_{k=1}^{K} \tilde{\gamma}_k^{s+} \Delta R_{s,t-k}^{+} + \\
& \sum\nolimits_{i=0}^{I} \tilde{\beta}_i^{s-} \Delta W_{t-i}^{-} + \sum\nolimits_{k=1}^{K} \tilde{\gamma}_k^{s-} \Delta R_{s,t-k}^{-} \\
& + \lambda_s \left(R_{s,t-1} - \tilde{\alpha}_s - \theta_s W_{t-1} - \tilde{\delta} t \right) + \upsilon_{s,t}.
\end{aligned}
\tag{2}
$$

In this ECM $\Delta W_{t-i}^{+}(\Delta W_{t-i}^{-})$ represents positive (negative) cost changes in period $t-i$ and $\Delta R_{s,t-k}^{+}$ $(\Delta R_{s,t-k}^{-})$ captures positive (negative) retail price changes in period $t-k$.[14]

This model is interesting for its own sake since it not only solves the co-integration problem between the retail price and the refinery price, but also has an interpretative interest. In particular, the coefficients $\tilde{\beta}_i^s$, $i = 1, \ldots, n$, represent the short-run response in the retail price to a change in the refinery price for gas station s, and the coefficients $\tilde{\gamma}_k^s$, $k = 1, \ldots, n$, capture the short-run response in the retail price for gas station s to its own price k periods back (persistence effect). The error-correction term $R_{s,t-1} - \tilde{\alpha}_s - \theta_s W_{t-1} - \tilde{\delta} t$ can be interpreted as the retail price deviation (one period lagged) from its long-run relationship with costs. We can think of the elements of the co-integration vector, $\tilde{\alpha}_s$ and θ_s as the coefficients arising from a regression of prices on contemporaneous costs, with $\tilde{\alpha}_s$ the long-run mark-up for station s and θ_s the long-run response in station s prices to cost changes. Additionally, λ_s measures the short-run correction in current prices that helps prices to return to their long-run equilibrium with costs. When prices exceed the cost by more than the long-run mark-up, a downward pressure on the price takes place until the long-run equilibrium is reached and the long-run mark-up is reestablished. The opposite occurs when the price is below the cost by more than the long-run mark-up. This means that we expect λ_s to be negative, and that the convergence towards the long-run equilibrium is faster the closer λ_s is to 1.

The main tool used in the rockets and feathers literature to illustrate the estimated amount of asymmetry is the difference in cumulative response functions (hereinafter CRF's). A CRF is the cumulative estimated price change for gas station s in period $t+j$ after a \$1 change in the cost variable in

[12] There could also be asymmetry in the speed of adjustment and volatility. Asymmetric price speed adjustment takes place when the number of periods that retail prices take to adjust to a long-run equilibrium after a cost increase is larger than that after a cost decrease, while asymmetric volatility takes place when positive shocks have a larger impact on volatility than negative shocks of the same size.

[13] In Tables A1 and A2 in the appendix we report the tests that show that the refinery and retail prices cointegrate and that the price series have unit roots.

[14] Formally, $\Delta W_{t-i}^{+} \equiv \max\{0, \Delta W_{t-i}\}$, $\Delta W_{t-i}^{-} \equiv \min\{0, \Delta W_{t-i}\}$, $\Delta R_{s,t-k}^{+} \equiv \max\{0, \Delta R_{s,t-k}\}$ and $\Delta R_{s,t-k}^{+} \equiv \min\{0, \Delta R_{s,t-k}\}$.

period t. In other words, the CRF in period $t+j$ is just the accumulated predicted price change up to period $t+i-1$ plus the predicted price change for period $t+j$. The CRF is a non-linear function of the estimated parameters in equation (2), given that a cost change in period t results in an adjustment in period $t+j$ that will be the sum of the estimated parameters and the error correction term during the n weeks that the adjustment to the long-run equilibrium takes place. When a cost increase takes place, the CRF function is denoted by $B^{s^+}_{t+j}$, while when a cost decrease takes place this is denoted by B^{s-}_{t+j}. In fact, it is easy to show that

$$B^{s^+}_{t+j} = B^{s^+}_{t+j-1} + \tilde{\beta}^{s+}_{j-1} + \sum\nolimits_{k=1}^{K} \tilde{\gamma}^{s+}_k \Delta R^+_{s,\,t+j-k} + \lambda_s \left(B^{s^+}_{t+j-1} - \theta_s \right)$$

and

$$B^{s^-}_{t+j} = B^{s^-}_{t+j-1} + \tilde{\beta}^{s-}_{j-1} + \sum\nolimits_{k=1}^{K} \tilde{\gamma}^{s-}_k \Delta R^-_{s,\,t+j-k} + \lambda_s \left(B^{s^-}_{t+j-1} - \theta_s \right)$$

Because the aim of this paper is to determine the existence of amount asymmetry in the market as a whole and not at a particular gas station, we seek to estimate $\left(E\left(\tilde{\beta}^{s+}_i\right), E\left(\tilde{\beta}^{s-}_i\right), E\left(\tilde{\gamma}^{s+}_k\right), E\left(\tilde{\gamma}^{s-}_k\right)\right)$ for $i = 0, \ldots, I$ and $k = 1, \ldots, K$; that is, the average across gas stations of adjustment and persistence coefficients; and $E\left(B^{s^+}_{t+j}\right)$ and $E\left(B^{s^-}_{t+j}\right)$ for $j = 0, \ldots, J$.

Assuming that these parameters are estimated correctly, the rockets and feathers hypothesis for a given market is then tested using two alternative null hypotheses:

H^1_0: the difference between the CRF for positive cost shocks and that for negative cost shocks for each period is nil, i.e., $E\left(B^{s^+}_{t+j}\right) - E\left(B^{s^-}_{t+j}\right) = 0$ for each $j = 1, \ldots, J$; and

H^2_0: the equality of the vector of coefficients for positive and negative cost shocks, i.e., $E\left(\tilde{\beta}^{s+}_i\right) = E\left(\tilde{\beta}^{s-}_i\right)$ for $i = 1, \ldots, I$.

Wald tests are used in both cases to test the symmetry hypothesis; the standard errors for the CRF's are obtained by means of the Delta method.

IV. THE DATA

Price data comes from a random sample of 50 gas stations (just over 10% of the population) located in Santiago, Chile, which is collected each Wednesday by the National Consumer Protection Agency. The price data has a weekly frequency and the period considered corresponds to the first

TABLE I
DESCRIPTIVE STATISTICS

Variable	Mean	Std. Dev.	Min.	Max.	N
Retail price	216.9	38.13	146.57	321.45	7964
Refinery price	173.9	33.49	107.60	264.58	181
Change in retail price	0.55	10.11	– 32.41	23.95	7920
Increases in retail price	8.13	5.99	0.04	23.95	4008
Decreases in retail price	– 8.14	7.06	– 32.41	– 0.06	3466
Change in refinery price	0.46	9.54	– 30.62	20.19	180
Increases refinery price	8.24	4.97	1.24	20.19	84
Decreases refinery price	– 9.36	6.81	– 30.62	– 1.67	65

week of March, 2001, up to the second week of August, 2004. Price information is collected directly from gas station signboards and not by interviewing gas station managers. This provides us with panel data with a time dimension equal to 181 periods and a cross-section dimension equal to 50 gas stations for a total of 9,050 data points. The survey also provides information concerning the specific location of each station and the brand identity under which the gasoline is sold. However, the sample is restricted to the 44 gas stations from which information for the whole period is considered. This means that the sample used contains 7,964 observations.

The paper focuses on 93 octane gasoline since this is the most important among the three different types of gasoline sold in Santiago, and studies the relationship between this type of gasoline and the price at the Concón refinery. Both of them are net of taxes and of compensations made by the oil stabilization fund.

Table I offers descriptive statistics in nominal Chilean pesos for the price series studied.[15] The average retail price over the whole period was $216.9 (U.S. $0.41) per liter and the average refinery price was $173.9 (U.S. $0.33). The table also shows that considering positive and negative price variations, on average the price increased $0.55 (U.S. $0.001) per liter during the whole period. On average the retail price rose 91 times and this meant an average increase of $8.12 (U.S. $0.015) per liter, while on average it decreased 79 times for an average decrease of -$8.13 (U.S. $0.015) per liter.

The average refinery price changed was $0.46 per liter. There were 84 positive changes and 65 negative ones. Thus, the refinery price did not vary in 31 opportunities. Price increases resulted in an average increase of $8.23 (U.S. $0.015) per liter and price falls meant an average decrease of -$9.36 (U.S. $0.017) per liter.

Figure 1 shows the retail and refinery price series for 93-octane gasoline. Their behavior is quite similar and in fact the correlation coefficient is 98.3%, which suggests that there is little delay in passing price decreases and price increases–that is, the adjustment process is quite symmetric. In

[15] One American dollar is equivalent to 530 Chilean pesos. Thus, the average retail price per liter is net of taxes equal to U.S. $0.41 and the refinery price is U.S. $0.33.

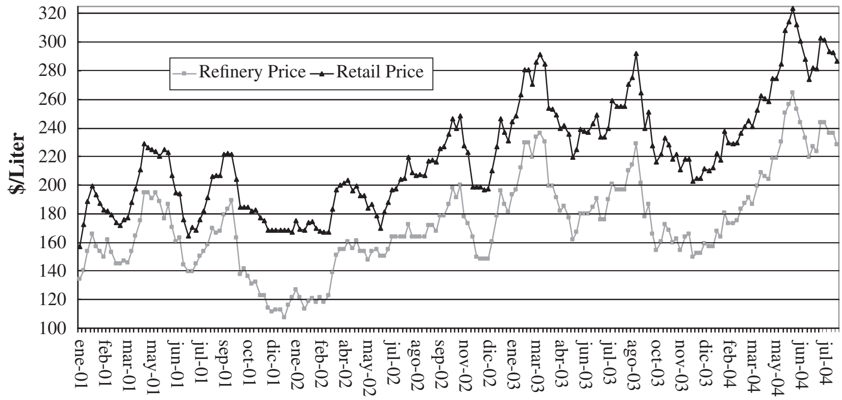

Figure 1
Chilean gasoline prices January, 2001–August, 2004.

addition, the null hypothesis that the variance of the refinery price series and that of the retail price series cannot be rejected at 5% significance level.[16]

V. RESULTS

Table II summarizes all results.[17] We present two alternative estimates. The first column of Table 2 shows the random effects (hereafter RE) estimates, which are valid under parameter homogeneity and no cross-section dependence,[18] and in the second column we present a random coefficient estimator that takes into account cross-section heterogeneity and cross-section dependence (hereafter RC-CCE). This amounts to including lagged cross-sectional averages of the retail prices in the regressions, which is estimated using the well-known Swamy random coefficient estimator.[19]

The reason for focusing on an estimate different from the random effect is that it is likely that the actual marginal cost faced by a station varies across stations; it is also likely that gas stations face common shocks that are not

[16] The 95 and 97-octane gasoline show the same behavior. That is, the hypothesis that the variance of different price series are equal cannot be rejected.

[17] Lag-lengths of the changes in the cost variable were identified using a general-to-specific technique starting from a maximum number of 6 lags. The number of lags for the dependent variable was chosen to ensure that there is no serial correlation.

[18] The random effect estimator is chosen over the commonly used fixed-effect estimator since the Hausman test for fixed versus random effects, which tests the null hypothesis that $E(\upsilon_{st}|X_{st}) = 0$, cannot reject this hypothesis with any level of significance (the Hausman statistic is 10.39 and the p-value is 1).

[19] This estimator was proposed by Pesaran [2005b] to deal with cross-section dependence and parameter heterogeneity. It is based on the assumption that the disturbances are assumed to contain one or more unobserved (latent) factors which may influence each unit differently.

TABLE II
REGRESSION RESULTS FOR ERROR CORRECTION MODEL (DEPENDENT VARIABLE $\Delta R_{s,t} = R_{s,t} - R_{s,t-1}$)

Variable	Random Effects (RE) Coefficient (Std. err)	Random Coeffieccient and Cross-sections Dependance (RC-CCE) Coefficient (Std. err)
W_{t-1}	0.084 (0.017)**	0.228 (0.013)**
R_{t-1}	– 0.101 (0.020)**	– 0.261 (0.015)**
t	0.017 (0.004)**	0.047 (0.003)**
ΔW_t^+	1.063 (0.009)**	1.062 (0.007)**
ΔW_{t-1}^+	0.146 (0.024)**	– 0.002 (0.023)
ΔW_{t-2}^+	0.141 (0.016)**	0.104 (0.018)**
ΔW_{t-3}^+	0.010 (0.017)	– 0.044 (0.023)
ΔW_{t-4}^+	– 0.051 (0.033)	– 0.108 (0.021)**
ΔW_{t-5}^+	0.039 (0.033)	0.014 (0.020)
ΔW_{t-6}^+	– 0.109 (0.017)**	0.102 (0.020)**
ΔW_t^-	0.901 (0.005)**	0.895 (0.007)**
ΔW_{t-1}^-	0.158 (0.026)**	– 0.068 (0.030)*
ΔW_{t-2}^-	0.046 (0.025)	– 0.158 (0.026)**
ΔW_{t-3}^-	– 0.057 (0.020)**	– 0.171 (0.024)**
ΔW_{t-4}^-	0.149 (0.025)**	0.055 (0.025)*
ΔW_{t-5}^-	– 0.027 (0.020)	– 0.150 (0.023)**
ΔW_{t-6}^-	0.108 (0.023)**	0.033 (0.022)
ΔR_{t-1}^+	– 0.129 (0.020)**	0.041 (0.021)
ΔR_{t-2}^+	– 0.107 (0.015)**	– 0.048 (0.017)**
ΔR_{t-3}^+	0.023 (0.017)	0.091 (0.021)**
ΔR_{t-4}^+	0.063 (0.036)	0.139 (0.019)**
ΔR_{t-5}^+	0.047 (0.032)	0.088 (0.018)**
ΔR_{t-6}^+	0.073 (0.017)**	0.084 (0.018)**
ΔR_{t-1}^-	– 0.123 (0.026)**	0.111 (0.030)**
ΔR_{t-2}^-	– 0.061 (0.025)*	0.135 (0.027)**
ΔR_{t-3}^-	0.057 (0.022)**	0.170 (0.024)**
ΔR_{t-4}^-	– 0.146 (0.025)**	– 0.052 (0.025)*
ΔR_{t-5}^-	0.010 (0.022)	0.137 (0.023)**
ΔR_{t-6}^-	– 0.100 (0.022)**	– 0.030 (0.022)
$\frac{1}{T}\sum R_t^+$		– 0.836 (0.067)**
$\frac{1}{T}\sum R_t^-$		– 1.100 (0.063)**
$\frac{1}{T}\sum R_{t-1}$		– 0.070 (0.046)
Constant	4.657 (0.982)**	11.208 (0.597)**
No of observations	7656	7656
No of stations	44	44
R^2	0.89	
$E\theta_s$	0.83	0.87
F-Test (p-value)	36.16 (0.00)	
Hausman (RC vs RE) (p-value)		58.01 (0.003)
Swamy Test (p-value)		1261.80 (1.00)
LM-test (p-value)	66541.42 (0.00)	
Estimated ρ (p-value)	0.01 (0.63)	
Serial Correlation (p-value)	0.23 (0.63)	

Robust standard errors in parenthesis.
**significant at 1% level;
*significant at 5% level.

entirely captured by shocks to the refinery price, or there are omitted (unobservable) global effects. If no correction for cross-section dependence is provided, then the standard estimation methods, as well as those dealing with heterogeneity across cross-section units, will be inconsistent if the global shocks are correlated with the regressors. The consequences of ignoring cross-section dependence may imply that pooling provides little

TABLE III
WALD TESTS BY ESTIMATION METHOD (p-VALUES IN PARENTHESIS)

H_0	RE	RC-CCE
$E_s\tilde{\beta}_i^{s+} = E_s\tilde{\beta}_i^{s-}$	222.28	364.05
	(0.00)	(0.00)
$E_s\theta^s = 1$	397.52	312.69
	(0.00)	(0.00)

gain in precision over single equation estimation. Nonetheless, this is an empirical question that requires empirical testing.

The Breusch-Pagan LM test overwhelmingly rejects the null of cross-section independence, and the parameter-homogeneity hypothesis, tested using a Hausman test, which is the procedure suggested by Im, Pesaran and Smith [1996], overwhelmingly rejects the null of homogeneity (p-value = 0.005).[20] Thus, according to these tests, the statistically correct estimator is the one that takes into account cross-section dependence and parameter heterogeneity.

Despite the possible differences between the two estimators, Table III shows that under each estimator the null hypothesis H_0^2 is rejected at a 5% significance level, and the long-run coefficient is almost identical and statistically different from 1 in each case, also at a 5% significance level. That is, no matter what estimator we use, the rockets and feathers phenomenon as measured by hypothesis H_0^2 is present in the local gasoline market studied here.

The first hypothesis (i.e., H_0^1) is studied in Figure 2 by mean of comparing the shape of the CRF's. Figure 2 shows the CRF's for the RC-CCE estimator, while Figure 3 shows the difference between CRF's for this estimator. In each figure the dotted lines represent 95% confidence bounds.[21]

The darker line in Figure 2 represents the estimated retail price response (in pesos per liter) to a one-time one-peso per liter increase in the refinery price. Thus, a one peso increase in the refinery price leads to a 1.062 peso increase in the retail price during the first week, then a small decrease occurs the following week, with a maximum increase of 1.162 after five weeks followed by a permanent decrease towards the long-run equilibrium. The starred-lighter line is the estimated retail price response to a one-time one-peso per liter decrease in the refinery price. Thus, a one peso decrease in refinery price leads to a 0.895 peso decrease in the first week, then a further

[20] Note, however, that the Hausman test is not a test for heterogeneity itself, rather, it is a test for the consequences of such heterogeneity on the consistency of different estimators. However, as seen in Table 2 the Swamy test also rejects the null of parameter homogeneity.

[21] The confidence bounds are derived using the delta method and imposing constraints on the intertemporal covariance structure.

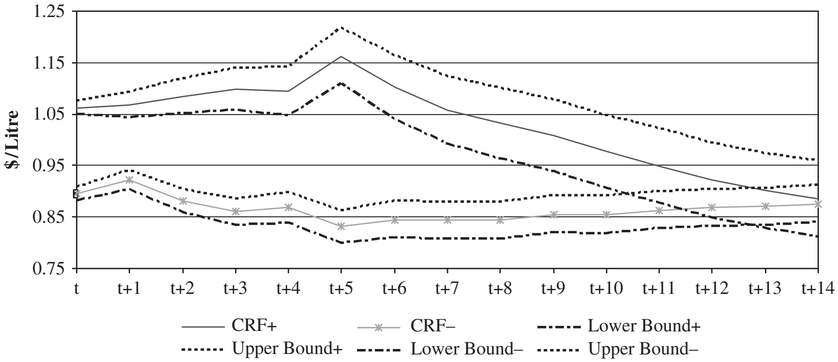

Figure 2
Cumulative response functions for positive and negative shocks: cross-section dependence and cross-section heterogeneity corrected.

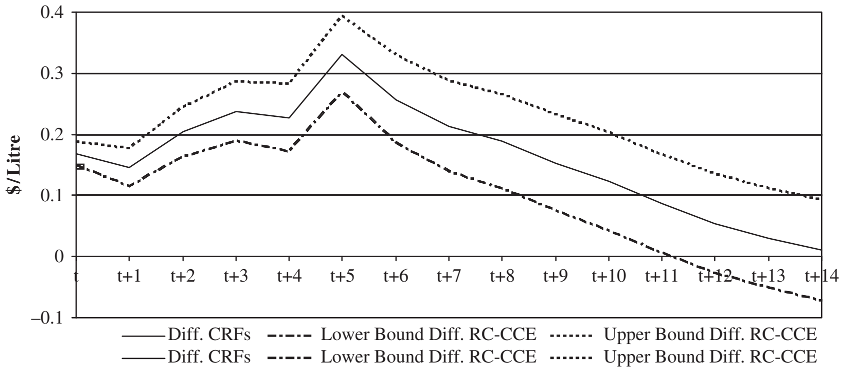

Figure 3
Difference in cumulative response functions: cross-section dependence and cross-section heterogeneity corrected.

decrease of almost 3 cents takes place, and subsequently a relatively smooth adjustment towards the long-run equilibrium is observed. It is relatively clear from the graph that the increases are passed along faster than the decreases and that the long-run equilibrium is reached after 11 weeks. In fact, from the difference in the CRF's that is shown in Figure 3, one can see that the difference in the CRF's is different from 0 with a 95% confidence only up to the eleventh week. This confirms the asymmetric adjustment pattern found by the Wald test applied to the null hypothesis H_0^2 shown in Table III.

Regarding the estimation of the long-run relationship between retail prices and the refinery price, we found that the coefficient on lagged costs in the error correction term θ_s is 0.87 and is statistically different from one.[22] This may seem counter-intuitive. However, a plausible explanation of $\theta_s < 1$ is that each station changes its price by a constant fraction of its cost change. Bulow and Pfleiderer [1983] show that for a residual demand curve of the form $p = a - bq^{\delta}$, prices will rise always by a lower amount than the rise in costs. The reasoning behind this is that a firm always changes its price by a fraction of its cost change, but it does not pass on the entire cost change. In fact, the fraction of the cost variation being passed on is a constant equal to $\frac{1}{\delta+1}$, which is independent of the elasticity of residual demand at any given point. While this result is derived from a highly simplified model, if gas stations in our data set face a residual demand, such as that specified above with a coefficient δ equal to 0.149, the implied fraction of the cost being passed on is 0.87.

While the Wald test and the CRF's show that there is pattern asymmetry, the consequences for consumers are difficult to ascertain from just looking at the CRF's. Borenstein *et al.* [1997] have proposed examining the gain for consumers for a \$1 increase in the refinery price over the lifetime of the price adjustment, in relation to the loss to consumers over the adjustment process from an equal size decrease in refinery price. Integrating the difference between the two CRF's over the whole adjustment process yields an estimate of the cost faced by consumers due to the asymmetric adjustment pattern:

$$\Delta\,\text{Consumer Cost} = \Delta C_n = \int_{j=1}^{n} \left(E_s B_{t+j}^{s^+} - E_s B_{t+j}^{s^-} \right),$$

where $E_s B_{t+j}^{s^+}$ and $E_s B_{t+j}^{s^-}$ are the CRF's for positive and negative cost shocks, respectively. Under linear interpolation between points, ΔC_n is the difference between the two CRF's in Figure 3 from week t to week $t+20$.

Figure 4 presents the estimated ΔC_n and its 95% confidence bounds. This indicates that the total cost asymmetry increases up to week $t+15$, four weeks after the long-run equilibrium is reached. The cost is significantly different from 0 in each week and it reaches a maximum of \$2.31 pesos (U.S. \$0.0044) per liter when a one time one peso change in the refinery price takes place. The extra cost during the whole adjustment period of a one time \$10 (U.S. \$0.18) per liter increase is \$5.3 (U.S. \$0.01 or 3.5%) per liter, while that due to a one time \$10 per liter decrease is \$17.9 (U.S. \$0.034 or 11.9%) per liter. For an individual whose weekly consumption is 40 liters, this means that the rockets phenomenon results in an extra cost of \$213.9 (U.S. \$0.40 or) and the feathers phenomena in an extra cost of \$712.9 (U.S. \$1.35). In other words, the rockets phenomenon results in an extra cost of 4.1% and the feathers phenomenon in an extra cost of 13.7% over the whole adjustment period.

[22] Lewis [2007] and Verlinda [2008] find that this is statistically greater than 1.

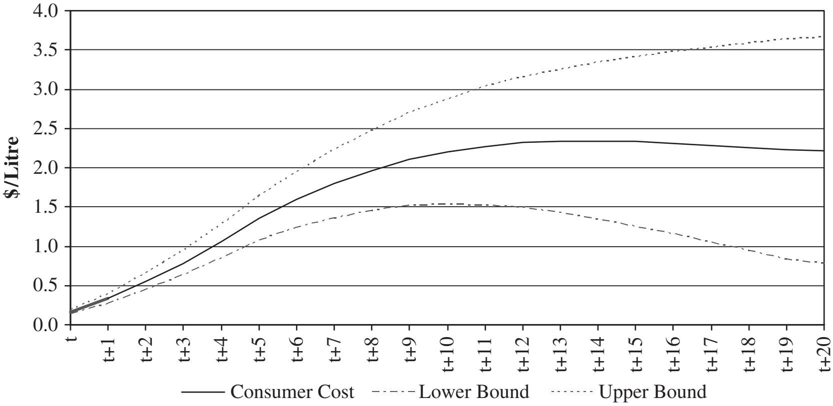

Figure 4
Estimated consumer cost of a one time $1 per liter price change: cross-section dependence and cross-section heterogeneity corrected.

To end this section, we shall illustrate the consequences of ignoring cross section heterogeneity and cross section dependence. Because of the large number of parameters estimated, this is hard to do by just comparing the estimated coefficients for each estimator. So, we will proceed by constructing the CRF's for the RE estimator; that is, when cross-section dependence and cross-section heterogeneity are ignored, and compare them with CRF's for the RC-CCE estimator.

Figure 5 compares the asymmetric pattern measured by the difference in CRF's for the RC-CCE estimator (starred-lighter line) with that difference for the RE estimator (darker line). Again, dotted lines represent 95% confidence bounds.

When parameter heterogeneity and cross-section dependence are ignored, the pattern asymmetry is quite different compared to when these two problems are accounted for. First, the RE estimator predicts that the long-run equilibrium is not reached after 14 weeks from the period in which the initial shock took place, while the RC-CCE estimator predicts that it is reached within 11 weeks. Second, the RE estimator sub-estimates the asymmetry during the first 10 weeks and overestimates it from the eleventh week onwards.[23]

It is clear from the analysis that ignoring cross section dependence and cross section heterogeneity results in a bias in the asymmetric price adjustment pattern. However, for the local gasoline market studied here, the

[23] Additional estimates made suggest that ignoring cross-section dependence results in a downward bias of the asymmetric adjustment price for the whole period.

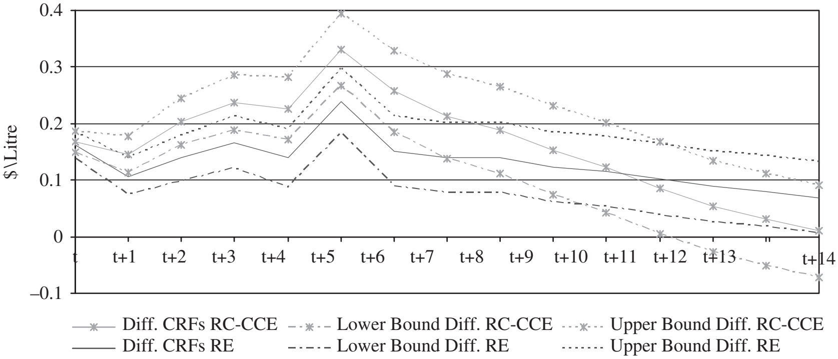

Figure 5
Difference in cumulative response function by estimation method: cross-section dependence and cross-section heterogeneity corrected vs. random effects.

bias is not so significant as to provide a different answer to the question of whether the price adjustment process is symmetric or not. This, however, may not to be true for all local markets.

VI. LOCAL MARKET POWER AND PRICE ASYMMETRY

Leaving aside the issues of cross section dependence and parameter heterogeneity, the most straight-forward approach to study the effect of market power on price adjustment asymmetry is the procedure adopted by Deltas [2008], which interacts the exogenous variables with observed margins in the ECM. He examines monthly average retail price data at the state level and finds that states in which average margins are higher show a more asymmetric adjustment pattern. The validity of this approach rests heavily on the assumption that the average margin in a state reflects the degree of local market power at retail level. To the extent that the characteristics of each state's gasoline market are similar, the average cost across states will vary across them as a function only of state specific wholesale spot prices, which are readily observable to the author.

At station level, however, the difference between the retail price and the refinery price is not necessarily the true margin. Indeed, the actual marginal cost faced by a given station varies across stations for reasons such as the difference in the type of contracts under which stations are operated. This makes the approach adopted by Deltas [2008] somewhat inadequate, as the result could be driven by a spurious relationship between the average margins and price-asymmetry. For instance, suppose that stations that operate under brand 1 have a lower margin than unbranded stations. Based on the Lerner index of market power, this would lead to the conclusion that brand 1 stations have lower local market power than unbranded stations. This would be wrong if brand 1 stations are operated under a different contract which entails a lower marginal cost than that for unbranded stations; although brand 1 stations charge on average a lower price, they do not necessarily have lower margins than unbranded stations, and they could be even greater.

Rather than interacting margins with the regressors in the ECM studied here, we contrast the asymmetry across different brands by means of comparing the CRF's for each brand, and between stations with higher than average margins and stations with lower than average margins. In particular, the statistical exercise carried out in this section is as follows. We estimated the ECM in equation (2) station-by-station, augmented with cross-section averages of the dependent variable, and then obtained the weighted average of the parameters by Brand and Margin. In other words, we obtained the RC-CCE estimator by Brand and Margin. In each case, we use this estimator to test H_0^2 and construct the corresponding CRF's.

Because the average margin over the sample period for each station used here, which is based on the difference between the retail price and the refinery price in each period, may not be the true margin, the results regarding the comparison between firms with high and low margins must be treated with caution. Nonetheless, the fact that margins are not interacted with regressors ensures that the coefficients used to construct the CRF's are the correct ones in the sense that they account for parameter heterogeneity and cross-section dependence. This implies that if the ranking of gas stations by true margins is preserved according to the margin calculated here, the exercise undertaken provides the right answer.

Brand Identity. There are five retail distribution companies in Santiago. Two of them (YPF and JLC) are pooled together under the unbranded label. Table IV reports the average retail price, the average margin by brand and the number of gas stations by brand. Copec has the largest market share with 51% of the market and the unbranded stations have the smallest market share, which is 10%. It is easy to observe in the table that both the average price and the average margin for branded gas stations are quite similar to one another, but greater than those for unbranded gas stations.

In Table V, Wald tests for H_0^2 and the null hypothesis that the long-run coefficient is equal to one are presented. It readily follows from the table that

TABLE IV
AVERAGE PRICE AND MARGIN BY BRAND (NUMBER OF STATIONS)

Brand	Copec (19)	Shell (12)	Esso (7)	Unbranded (6)	Total
Price	217.5	216.9	216.7	215.8	216,9
Margin	43.6	43.0	42.7	41.8	42.9

TABLE V
WALD TESTS BY BRAND (p-VALUES IN PARENTHESIS)

H_0	Copec (19)	Shell (12)	Esso (7)	Unbranded (6)
$E_s\tilde{\beta}_i^{s+} = E_s\tilde{\beta}_i^{s-}$	171.34	118.42	64.49	53.18
	(0.00)	(0.00)	(0.00)	(0.00)
$E_s\theta^s = 1$	125.87	101.38	42.57	52.98
	(0.00)	(0.00)	(0.00)	(0.00)

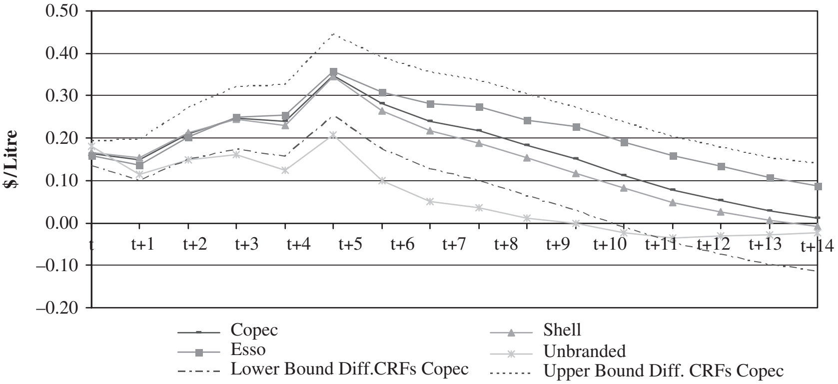

Figure 6
Difference in cumulative response functions for branded and unbranded gas stations: cross-section dependence and cross-section heterogeneity corrected.

the two null hypotheses are rejected at a 5% significance level for both branded as well as unbranded gas stations.

The fact that the null is rejected in each case does not mean that the adjustment process is the same for each brand. To study the difference across brands, the tabular approach is again not very useful and thus the CRF's for each brand are obtained.

Figure 6 shows the difference between the CRF's for each brand – that is, $E_sB_{t+j}^{s^+} - E_sB_{t+j}^{s^-}$ where the expectation is taken over all gas stations belonging to the corresponding brand. The first thing to notice is that the asymmetry during the first week, measured by $E_sB_t^{s^+} - E_sB_t^{s^-}$ is almost identical for branded as well as unbranded stations. Second, the asymmetry for the whole period is quite similar for the branded stations Esso, Copec and Shell, while it is different from that for the unbranded stations. In fact, the asymmetry for the highest priced brand (Copec – the darkest line) is statistically different from the asymmetry for the unbranded stations (the starred-lighter line) between the second and the eleventh week, while it is statistically the same for the three branded stations. The estimated asymmetry peaks after five weeks for Copec at 35 cents per liter, while that for unbranded stations also peaks after five weeks, but only reaches 20.7

cents per liter. Thus, the difference between these two groups is estimated at nearly 15 cents per liter.

If we interpret the asymmetry as a cost that consumers bear from buying a particular brand, then the result suggests that buying from branded stations results in an extra cost, which is statistically significant, in relation to buying from an unbranded station. Moreover, as also concluded by Verlinda [2008], if the underlying process generating the asymmetry is implicitly collusive, then this suggests that in the event of a shock collusion it is broken down more rapidly for unbranded stations than for branded stations.[24] This is consistent with the idea that collusion is easier to sustain the more important is horizontal differentiation, where this is understood as brand loyalty.

High and Low Margin Gas Stations. Next we look at the relationship between market power measured by average margins and the price-adjustment pattern. The sample is split between high-margin stations and low-margin stations, where a station is classified as high margin when its average margin over the sample period is greater than the average margin over all stations, where the average margin in station s over the sample period is $M_s \equiv \sum_{t=1}^{T} \left(R_{s,t} - W_t\right)/T$ and the average margin over all stations is $M \equiv {}_{s=1}^{S} M_s/S$. Thus, if $M_s > M$, then station s is classified as a high-margin station, while if the opposite occurs, station s is classified as a low-margin station.

The distribution of gas stations between the two groups is as follows. There are 30 gas stations that are classified as high-margin stations and 14 as low-margin stations. Of the 19 gas stations operated by Copec, 79% are classified as high-margin stations; 75% of the 11 stations operating under the Shell brand are in this group; Esso has 7 gas stations in the sample, of which 43% (3 stations) are classified as high-margin stations; and 4 of the 6 unbranded stations belong to the group made up of low-margin stations.

In Table VI, Wald tests for H_0^2 and the null hypothesis that the long-run coefficient is equal to one are presented. The results show that both high and low-margin stations present an asymmetric price adjustment; that is, the null hypothesis H_0^2 is rejected at a 5% significance level in each case, and that the long-run coefficient is smaller than one in both cases, also at a 5% significance level.

In Figure 7 we look at the issue of local market power by examining the asymmetry between low (starred-lighter line) and high-margin stations (darker line), measured again by the difference between the CRF's. The

[24] Horizontal differentiation on the one hand limits the short-run gains from undercutting rivals, since it is more difficult poaching consumers who are loyal to other brands, but on the other hand decreases the punishment from deviation since it limits the severity of price wars. When the gain from deviations falls more than the loss from punishment as horizontal differentiation rises, collusion will be easier to sustain.

TABLE VI
WALD TESTS BY MARGIN (p-VALUES IN PARENTHESIS)

H_0	High Margin (30)	Low Margin (14)
$E_s\tilde{\beta}_i^{s+} = E_s\tilde{\beta}_i^{s-}$	233.85	156.25
	(0.00)	(0.00)
$E_s\theta^s = 1$	232.74	100.80
	(0.00)	(0.00)

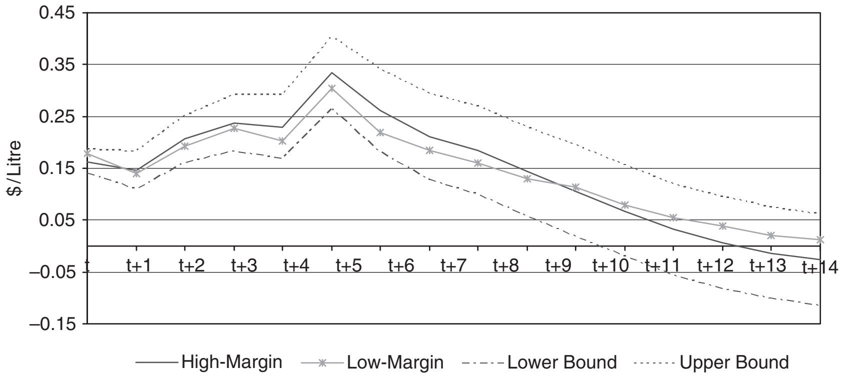

Figure 7
Difference in cumulative response functions by margin: cross-section dependence and cross-section heterogeneity corrected.

figure reveals that in the first week asymmetry is slightly higher for low-margin stations, and then the opposite occurs from the second week onwards until the tenth week, which is the point when this is reversed. The maximum difference is reached six weeks after the initial shock, a period in which a 4 cents per liter difference is reached. However, the difference in CRF's for low-margin stations is always contained within the confidence bounds for the difference in CRF's for high-margin stations, which means that the difference between these two types of station is not statistically significant. Thus, if the ranking of stations in terms of their average margin, based on our measure of margin, maintains the ranking of stations based on the true average margin, then local market power has no statistically significant effect on the asymmetric price adjustment process found in the data.[25] This result could be explained by the fact that the characteristics of the Chilean market result in wholesale and retail companies' gasoline, transportation and storage costs being almost if not identical, which together with the

[25] This conclusion tends to be confirmed when one makes a pair-wise comparison between the price asymmetry across stations. However, there are statistically significant differences in the price adjustment patterns between the station with the lowest margin and that with the highest margin.

announcement policy facilitates the adoption of a collusive pricing policy in which most gas stations follow ENAP's recommendations very closely.

VII. CONCLUSIONS

Using a highly detailed station level data set, we have documented in this paper the existence of an asymmetric price behavior in the Santiago gasoline retail market. As in most developed countries, retail prices rise faster for a cost increase than they fall for a comparable cost decrease. This result is obtained after accounting for cross-section heterogeneity and cross-section dependence (problems largely ignored in the test for the existence of the rockets and feathers phenomenon). In addition, it is shown that price-response asymmetry is a general feature of the data. Neither brand identity nor local market power measured by average margins can fully explain the asymmetric adjustment price behavior found in our data, yet brand identity contributes measurably to its existence.

A particular feature of the rockets and feathers phenomenon in the Chilean market is the presence of large initial responses to cost shocks relative to the initial responses found in developed countries. However, this particular feature does not result in more symmetric adjustments and lower costs for final consumers relative to those in developed countries. In fact, the asymmetries found are in accordance with those found by Lewis [2007] and Verlinda [2008], which are two of the few studies that use gas station level data. A plausible explanation for this is the fact that ENAP announces each week by how much retail prices should vary if the refinery price change were fully passed on to final consumers. Thus, a policy of announcing price changes may not result in the intended result of ameliorating the rockets and feathers phenomenon, and furthermore the evidence suggests that this might be facilitating collusion since the policy of announcing price changes provides a natural focal point, hinders price wars, and increases the frequency of price adjustments.

While the data available does not allow us to investigate in more detail the influence of local market power on price asymmetry, if the price asymmetry found is the result of collusion, then the evidence here suggests that collusion is easier to sustain in the presence of differentiated products and homogeneity in cost structures across gas stations.

APPENDIX: UNIT ROOT TESTS AND COINTEGRATION TESTS

Table A1 shows the unit root tests for the panel data retail price series. The PANIC test is employed together with three more tests that allow for cross-section dependence to be reported. Three of the four tests reject the null hypothesis of a unit root. This suggests that cross-section dependence drives the data generating process, which leads to a consistent diagnosis of the integration properties of the price series and emphasizes the

TABLE A1
UNIT ROOT TESTS

Unit Root Tests*	Refinery Price	Critical Value (5%)
ADF(1)	– 2.06	– 2.88
PP	– 2.05	– 2.88
	Retail Price	Critical Value (5%)
O'Connell	– 23.61	– 9.74
Breitung and Das	– 1.76	– 1.71
Pesaran - CADF (t-bar)	– 7.27	– 2.12
Bai and Ng (PANIC) (DFFc(m))	– 2.57	– 2.86

*Regressions include an interceptPanel data H_o: all 44 timeseries in the panel are I(1) processes.

TABLE A2
COINTEGRATION TESTS

Cointegration Tests	Refinery Price	Critical Value (5%)
Engle-Granger*	– 4.55	– 3.37
	Retail Price	Critical Value (5%)
DF rho-statistic	– 42.77	– 1.96
DF t-statistic	183.98	1.96
Nharvey**	24.36	10.03

*H_o: Non cointegration,
**H_o: 0 common trends among the 44 series in the panel.

role of cross-unit co-integration in determining the presence of the rockets and feathers phenomenon.

Table A2 shows the results for the panel time series co-integration tests. The two versions of the KAO test, as well as the Harvey's test, reject the null hypothesis of no co-integration.[26]

These findings lead to the conclusion that the price series are integrated by of order one and that prices and costs are co-integrated, thus warranting the use of an ECM.

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[26] As panel tests are prone to over-rejection in the presence of cross-section co-integration, this result is not surprising.

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