# Testing Optimal Punishment Mechanisms under Price Regulation: the Case of the Retail Market for Gasoline

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# Testing Optimal Punishment Mechanisms under Price Regulation: the Case of the Retail Market for Gasoline

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**ABSTRACT**

We analyse the effects of a price floor on price wars (or deep price cuts) in the retail market for gasoline. Bertrand supergame oligopoly models predict that price wars should last longer in the presence of price floors. In 1996, the introduction of a price floor in the Quebec retail market for gasoline serves as a natural experiment with which to test this prediction. We use a Markov Switching Model with two latent states to simultaneously identify the periods of price-collusion/price-war and estimate the parameters characterizing each state. Results support the prediction that price floors reduce the intensity of price wars but increase their expected duration.

**Keywords** : price regulation, oligopoly supergame, Markov switching model, gasoline prices.

**JEL codes**: L13, L81, C32.

## 1. Introduction

The behaviour of retail gasoline prices has long been and still is the object of fierce public debate. Given its importance for the consumers and the apparently "suspect" behaviour of the integrated oil companies (the "majors"), many jurisdictions have regulated aspects of gasoline retailing. For example, refiners are forbidden to operate retail outlets in some U.S. states. Different types of price regulations are also enforced in several U.S. states and Canadian provinces. In this paper, we provide evidence of the effects of price floor regulation on the behaviour of retail gasoline prices in Montreal, the largest market in the province of Quebec.

The May 4, 1992 edition of Bloomberg's Oil Buyers Guide reported the conclusions of market studies indicating that "major Canadian oil companies were going to use price wars, new credit terms, and the strategic closure of service stations and refineries to squeeze independent gasoline retailers out of the market in central Canada".[1] A few years later, following a severe price war in the summer of 1996, the Quebec provincial government responded to the lobbying of independent gasoline retailers by establishing a price floor in December 1996. The regulation was motivated by the claim that price wars formed as a discipline device for the implementation of anticompetitive strategies. This was made very clear in a report by the local association of independent gasoline retailers: "The summer 1996 episode was very harmful for suppliers in Quebec. The price war, triggered by an integrated major, resulted in retail prices that were observed well below wholesale prices. It was so severe as to force several independent retailers either to close down temporarily or to exit the market" (pp. 7-8).[2]

By limiting the severity of price wars, the floor was thus rationalized as a mean to reduce the ability of firms to punish retailers deviating from a high price strategy. When the latter retailers were small independent suppliers, characterized by less favourable financial conditions than majors, the floor was viewed as a form of protection that would help to maintain competition in the market. The floor was introduced to limit the severity of price wars which, according to the independent retailers, were evidence of predatory behaviour by the integrated majors. The price floor was therefore viewed as a form of market protection for independent retailers which would thereby help to maintain competition in the market.

The floor is computed weekly and regionally as the sum of the wholesale (rack) price, transportation costs and taxes, and is the only type of economic regulation specific to Quebec's retail market for gasoline. The price floor computed in week *t* is the one which applies in the market during week *t*+1.

Our objective in this paper is to determine whether the effects of a price floor on retail gasoline markets is consistent with the predictions of a class of models which assume that there is no explicit collusion between firms. In a theoretical section, we review two types of such models which establish that a punishment price below a retailer's marginal cost makes economic sense, and that such severe price wars may be observed in the absence of regulation. A first group of models (Lambson, 1987, 1994; Häckner, 1996) relies on Bertrand oligopoly supergames in which, during collusive periods, firms price between the Nash equilibrium value of the stage game and the joint profit-maximizing price. If a firm deviates from the collusive strategy, all firms in the next period price below the one-shot symmetric Nash equilibrium price (a punishment), before returning to the collusive price to earn above normal profits. When deviation payoffs are relatively high, a price below marginal cost may be required in the punishment phase for tacit collusion to be sustainable. In this framework, by imposing a boundary on the set of available retail prices, a price floor restricts the severity of price wars. It can be viewed as a monkey wrench in the collusive works that contributes to making coordinated pricing strategies less stable. Theory predicts that a longer punishment period can compensate for a higher punishment price. This means that, when a price floor is introduced, the same collusive price can be sustained if the length of punishments is also a control variable of firms. However, the coordination mechanisms of these models lack realism in the sense that they make collusion sustainable with no punishment at equilibrium. But price wars <u>do</u> occur on real-world markets. In a second category of papers we consider (Porter 1983a, Green and Porter 1984, Slade 1989, 1992), unobserved random shocks in demand may induce price wars to appear in equilibrium. Taken together, these models lead to a simple prediction: the duration of price wars (spotted out as deep, irregular, unpredictable price cuts) should increase when a price floor is introduced. We believe the introduction of a price floor on retail gasoline markets in Quebec provides a natural experiment in which this prediction can be tested.

In this study, we do not attempt to integrate the many other models of price formation in gasoline markets. Their relevance to the present analysis of the impact of a price floor regulation is somewhat limited, as they do not allow for the possibility of pricing below marginal costs (not even out of equilibrium), nor rationalize the occurrence of price wars (in equilibrium). A prominent example of such a model is that of Rotemberg and Saloner (1986), who focus on the impact of observable independently and identically distributed changes in demand on the level of tacitly collusive prices. In a price-setting oligopoly supergame with infinite horizon and undifferentiated products, a deviation from the collusive path triggers a punishment mechanism, i.e. competitive pricing, in all subsequent periods.[3] The basic argument is that the incentives to deviate from collusion increase when demand is relatively high, since by undercutting its rivals a firm may capture a larger market share. Firms can sustain collusive profits in that case by lowering the collusive price and thereby making deviations less profitable. In equilibrium, we therefore only observe a "tuned" collusive price, which gets closer to the marginal cost during periods of high demand; it is not a punishment price. Haltiwanger and Harrington (1991) extend this model, assuming that demand moves cyclically in some deterministic way, and that firms' expectations of future demand change over time. Although their analysis confirms that the profitability of deviation from collusion is large when demand is high, periods of decline in demand are the most difficult times for firms to sustain collusion. A direct implication is that prices will be closest to the competitive level – and thus questionably interpreted as a price war – in early periods of recession, that is when demand is not only high but also falling. Again, these prices are charged above marginal costs – a feature that does not fit our empirical data in the episode that motivated the introduction of a price-floor – and remain of a tacitly collusive nature, in all equilibria.[4,5] The model is generalized by Bagwell and Staiger (1997), who assume that turning points are unpredictable. However, as noticed by the authors (see footnote 4, p. 82), the relevance of this assumption is questionable for the analysis of markets subject to seasonal demand fluctuations such as gasoline markets.

Finally, recent studies by Eckert (2002) and Noel (2003) also use price data for Canadian cities in order to explain features of retail gasoline prices. These two papers use the same alternating-move price-setting duopoly model, adapted from Maskin and Tirole (1988), in which a class of equilibria can be constructed that consists of price cycles. In these cycles, firms repeatedly undercut one another in the downward portion of the cycle. When the price reaches the competitive value, a war of attrition begins, until one firm raises its price back to the top of the cycle (i.e. above the monopoly price) and then the cycle is repeated. Eckert (2002) formulates an econometric version of the model of price cycles, and estimates it with daily data that describe prices in the city of Windsor (Ontario) from 1989 to 1994. He finds results consistent with the theoretical predictions that price reductions result when firms undercut rivals, and that the size of price restorations depends on the wholesale price. Noel (2003) develops an empirical framework to separate out different pricing phenomena, and applies it to the analysis of dynamic pricing behaviour in nineteen Canadian retail gasoline markets from 1989 to 1999. He finds that the most frequent pattern is one in which prices cycle rapidly, beginning with a large increase of short duration, before declining gradually over a longer time period. The evolution of prices in Montreal, in the period that precedes the 1996 price war, appears less easily classifiable in a given category of pricing phenomena than the patterns observed in other Canadian marketplaces.[6] In both papers, price increases are initiated when the retail price is close (but not *lower* than) the wholesale price, as suggested by the underlying theory.[7] By contrast, our reference to models of collusion, in which short price wars alternate with longer collusive episodes, is justified by the specific feature of the data we use, that is the occurrence of several and relatively large negative margins, in early 1995 and mid-1996.[8]

To test for collusive pricing and estimate the length of price wars, we use a Markov Switching Regression framework with two latent states to simultaneously identify the periods of price-collusion/price-war and estimate the parameters characterizing each state. Consistent with the theory, we allow regulation to influence both the state-conditional prices and the expected duration of each state. The switching regression is then estimated on weekly data for retail gasoline price margins in Montreal from 1994 to 2001. In summary, we find that the introduction of a price floor significantly increases margins during price wars, and also increases the persistence of those price wars. The net effect on average margins is near zero, as the increase in margins during price wars is offset by the rise in the average duration of price wars.

Section 2, below, motivates the specific prediction tested in this paper, by exploiting some insights from theoretical contributions on the possible impact of a price floor on firms' ability to sustain collusion. The empirical model and data are described in Section 3. The results are discussed in Section 4, and Section 5 presents concluding remarks.

## 2. Severe Price Wars and Binding Price Floors

It is well known that a price-setting oligopolist may find it profitable to undercut the price that maximizes industry profits. This may not be the case if deviations from the collusive price are credibly "punished" via lower industry prices for a period of time. When the punishment is sufficiently severe, the immediate gains to deviation are smaller than what is lost in the punishment phase. In a legal context which prohibits price cartels, price wars can thus be used as an enforcement device to maintain a tacit cooperative agreement. Accordingly, by imposing a constraint on firms' ability to dissuade deviations from a tacitly cooperative arrangement on the retail gasoline markets in Quebec, in 1996, the provincial government aimed at limiting the severity of price wars and firms' ability to sustain anti-competitive prices. Since then, the floor has been computed weekly and regionally as the sum of the rack price, transportation costs, and taxes. (We refer to it hereafter as retailer's marginal cost.) However, for this mechanism to be effective, it must be the case that a punishment price below a retailer's marginal cost makes economic sense, and that such severe price wars are observed in the absence of a price floor.

The literature on optimal penal codes in supergame oligopoly models provides a rationale for punishment prices which are below marginal cost. The specifications of these models, in their Bertrand version and with complete information, reflect the main properties of gasoline markets: sales are repeated, retailers use price as a strategic variable, prices are observable by all, rival sales can be monitored, price adjustment can be small and occur at low cost. In particular, Lambson (1987, 1994) characterizes optimal punishments for a large class of infinitely-repeated games, with price-setting sellers of a homogenous good and a constant marginal (and average) cost $c$.[9] Häckner (1996) builds on this by demonstrating that a symmetric two-phase "stick-and-carrot" structure is an optimal price path in a specific supergame duopoly model with differentiated products.[10] This means that, in a collusive period, all firms sell at a price that yields the highest sustainable level of profits.[11] This price, say $p^*$, is between the one-shot Nash equilibrium value $p^N$ and the joint profit-maximizing price $p^m$. When firms detect a deviation from the collusive strategy $p^*$, they both switch to a lower price $\underline{p}$ in a one-period severe punishment phase (the "stick"), where $\underline{p}$ is strictly less than $p^N$, before returning to $p^*$ to earn collusive profits again (the "carrot").[12] The higher the deviation payoffs, the lower the optimal punishment price must be for collusion to remain sustainable. This price may therefore need to fall below marginal cost $c$, in which case the price floor constraint $p \geq c$ would be binding.[13] When this occurs, one-period punishments are not sufficiently large to outweigh the gain from undercutting the collusive price. However, Häckner shows that collusion sustainability can be maintained by a $\tau$-period price war ($\tau > 1$), with a punishment price equal to $c$ in the first $\tau - 1$ periods, and to $p'$ in the $\tau$-th period, with $c \leq p' < p^*$. Then firms return to $p^*$ in the next period onward. By constraining the punishment price to remain above the marginal cost, it is possible that a prolonged price war can sustain the same collusive price as in the unrestricted case.[14]

Note that the relatively large number of geographically close outlets that are available to retail gasoline customers in an urban area contributes to a high level of product substitutability among consumers. In this theoretical framework, such substitutability implies larger market share gains (and related profits) for those who deviate from collusive pricing. Large incentives for deviations from collusive pricing in turn can call for a severe (below marginal cost) punishment price. This may motivate competition authorities to establish a price floor, as occurred in the province of Quebec.

As they can account for punishment prices below marginal costs, the above models shed some valuable light on the possible impact of a price floor on observed prices. However, these models lack realism in that they never produce price wars in equilibrium. This outcome is rooted in the simplifying assumption of complete information. On the grounds that price wars are frequently observed in real-world markets, Porter (1983a) and Green and Porter (1984) challenge the relevance of such models. They construct a dynamic model in which firms observe only their own production levels and a common market price. As the market demand has an unobservable stochastic component, firms face an inference problem. An unexpectedly low price may signal either a deviation from the collusive path by rival firms or a demand-reducing shock. In this set-up, firms produce a collusive total output until price drops below a trigger value, in which case they revert to the one-shot Nash equilibrium for a number of periods – the price war – before returning to the collusive strategy. Since this result is obtained in a quantity-setting oligopoly model, it should be invoked cautiously for the rationalization of observed price wars on gasoline markets.[15] For this reason, the proposition that unobserved random shocks in demand may induce price wars in equilibrium has been extended by Slade (1989, 1992) to a price-setting supergame model, which conforms more closely to the characteristics of a retail gasoline market. The main idea is that, although all prices are observable and rival sales can be monitored, when a shock has occurred firms do not know the true demand conditions unless they change their price and thereby precipitate a price war. In the model, price is measured above constant marginal costs, and pricing strategies are formulated as intertemporal continuous reaction functions which express a firm's price as a function of the previous-period prices charged by competitors.[16] A Nash equilibrium in prices can be computed for stationary demand. When an unanticipated change in demand occurs, a firm's price is no longer the best response to other firms' prices and a new equilibrium must be found. Each firm then uses the prices charged by competitors as signals to update its expectations about new demand conditions in a Bayesian fashion. Price wars are then rationalized as a learning process for parameter identification.[17]

By introducing an element of uncertainty, models with price wars in equilibrium gain in realism compared to studies with complete information. However, this comes at a cost, since the mentioned price-setting models in which price wars occur rule out pricing below marginal cost by construction.[18]

In the empirical work that follows, we draw on the lessons obtained from the two categories of models discussed above to investigate the effects of price floor regulation on the behaviour of retail gasoline prices in Montreal. In a nutshell, if punishment prices below marginal costs can be rationally justified, and if price wars can occur in the absence of complete information, we expect that the introduction of a price floor should lengthen the duration of price wars. This is the empirical prediction we test in what follows.

## 3. Econometric Implementation

### *Data and Variables*

Retail price data ($P_t$) have been provided by M.J. Ervin Inc., a Calgary-based firm which conducts a weekly survey on gasoline retail prices in all major Canadian markets. In each market surveyed, retail prices are collected by gasoline grade using a sample of self-service gas stations. Whenever possible, the same stations are surveyed each week. In the Montreal market, the survey covers approximately 20 stations. For our analysis, we use the average retail price for unleaded regular gasoline computed from all stations in the Montreal survey. The data cover the 1994-2001 period (416 weekly observations). Our analysis is limited to the price of unleaded regular gasoline as retail prices for all other grades follow closely unleaded regular gasoline prices.

Wholesale prices ($W_t$) have also been provided by M.J. Ervin. Those prices are "rack" prices (excluding taxes) posted everyday at wholesale distribution outlets. As for retail prices, we used the average (unweighted) weekly wholesale price computed from posted prices. No transportation costs are considered given the proximity of the retail market to the different wholesale distribution outlets.

In our empirical analysis, we use retail margins ($M_t$) rather than prices in order to eliminate effects coming from the wholesale market and thereby concentrate on retail market effects. Retail margins are computed as retail prices in week $t$ minus wholesale prices observed during the same week. Recall that the price floor for week $t$ is computed as the sum of the wholesale price observed during week $t$-1, corresponding taxes and transportation costs.

Finally, our empirical analysis includes a regulation dummy ($R_t$) equals to 0 until the price floor regulation was introduced during the last week of December 1996, and equal to 1 thereafter.

Figure 1 shows the evolution of retail margins over time. The pattern of fluctuations of the retail margins seems similar before and after the introduction of the price floor regulation; of course, large negative margins are no longer observed (as in early 1995 and mid-1996). (Small negative margins are still observed under regulation because the price floor is computed from the wholesale price observed the week before.) Deep price cuts can be observed in addition to variations of more limited magnitude, or cycles. These cuts are irregular, and seem unpredictable. At first glance, periods in which margins are either negative or very close to zero never last more than a few weeks. However, given the weekly nature of the data, each point of the graph may represent a constant price for up to seven days, implying that price wars may take place over tens of days. To see that, focus for example on the circumstances that motivated the introduction of a price floor, in mid-1996 about week 26 (when the margin is -5.875, the lowest value of our data). If a price war refers to circumstances in which the margin falls rapidly down to a level measured below a given threshold, say 1¢/litre, before returning to higher positive values, we obtain a three-week episode from week 25 to week 27, as made clear in Figure 2.[19] Another example can be observed in late-1996, shortly before the introduction of a price-floor, about week 37 (the margin is -1.1). If one keeps 1¢/litre as a working criterion, again one obtains a three-week price war, from 1996-36 to 1996-38.

**[Figures 1 and 2 here]**

Table 1 presents some descriptive statistics. We see that average margins are roughly the same before and after regulation but significantly less dispersed after regulation (the standard deviation is 2.61 before regulation and 2.00 after). Given that peaks are approximately of the same magnitude before and after regulation (see Figure 1) and that average margins are about the same size, small margins (i.e. deep price cuts) should be more prevalent after regulation, in accordance with theoretical predictions. This higher prevalence of small margins is not obvious in Figure 1, which calls for a more elaborate data analysis.

Toward this aim, in Table 2 we present the number of weeks (in %) where the margins are below and above average margins, before and after regulation. Before regulation, margins were below average 43% of the time while after regulation this number increased at 48%. Again, given that average margins were nearly the same before and after regulation, this result is also a sign (albeit a weak one) that periods of small margins are more prevalent since regulation. This provides some modest albeit encouraging indirect support for our theoretical prediction. We therefore now turn to a more formal statistical analysis of retail margins using a Markov Switching regression approach.

**[Tables 1 and 2 here]**

### *Econometric Model*

To test the prediction from Section 2, we need to allow for the structural relationships determining margins to vary depending on whether the industry is in a period of collusion or price war (competition). We further need to investigate how the duration of these different states may be influenced by changes in the regulatory environment. Estimation and inference is complicated by the fact that, while regulatory changes are directly observed, the presence or absence of price wars must be inferred indirectly.[20]

Of course, one could simply construct a binary variable to indicate which observations appear to correspond to price wars, then use traditional methods (e.g. Ordinary Least Squares) separately on the two distinct subsets of observations (for example, see Borenstein (1991, 1996)). While intuitive, this approach has serious problems. First, since the separation into price war/collusion is somewhat uncertain, some observations will be misclassified. That means this approach will produce biased and inconsistent estimates of the underlying relationships. Second, standard errors for the resulting regressions will ignore the contribution of uncertainty about the sample separation, making inference unreliable as well. The approach we adopt avoids both of these problems.

Porter (1983) and Lee and Porter (1984) address the same kind of problem using regime switching techniques, which estimate the structural parameters of each pricing regime together with probability that each observation may have been produced by a price war. We use an extension of their approach, inspired by Hamilton (1993)'s Markov Switching Models with time-varying transition probabilities.[21]

As our baseline model, we estimate the following system of equations by maximum likelihood

$$M_t = \alpha_i + \rho_i M_{t-1} + \beta_i R_t + \gamma_i W_t + \varepsilon_t \tag{1}$$

$$\Pr(S_t = i | S_{t-1} = i) = \Lambda(\varphi_i + \theta_i R_t + \omega_i W_t) \tag{2}$$

where $M_t$ , $W_t$ and $R_t$ are defined as earlier, $i$ = 1 for price wars and 0 otherwise, $S_t$ is the market state (1 for price wars and 0 otherwise) at time $t$, $\Lambda(.)$ is the logit cumulative distribution function, $\varepsilon_t$ is a an i.i.d. mean-zero normally-distributed error term with a standard deviation of $\sigma_\varepsilon$, and $\{\alpha_0,\alpha_1,\rho_0,\rho_1,\beta_0,\beta_1,\gamma_0,\gamma_1,\varphi_0,\varphi_1,\theta_0,\theta_1,\omega_0,\omega_1,\sigma_\varepsilon\}$ is the vector of unknown parameters to be estimated by maximum likelihood. We performed the usual diagnostics tests suggested by Hamilton (1996) for the fit of such models.

Since the two states in this model follow a first-order Markov chain, we can calculate the half-life of a regime (the length of time over which the probability of remaining in the same regime has fallen to 50%) as $\ln 0.5/\ln a$, where $a$ is the probability given by equation (2). Similarly, we can calculate the expected duration of the regime (in periods) as $1/(1-a)$. Furthermore, if $\{a_0,a_1\}$ are the probabilities of remaining in regimes 0 and 1 for one more period, then on average the market will spend fraction $(1-a_1)/(2-a_0-a_1)$ of the time in regime 0 (price collusion) and the remainder in regime 1 (price war).

The sign and significance of $\beta_i$ determine whether the introduction of a price floor has an impact on margins either during price wars ($i=1$) or collusion ($i=0$). Due to the dynamic nature of the model, the long-run impact of the regulation on margins in each regime will be $\beta_i/(1-\rho_i)$. The parameter $\theta_i$ tells us whether the price floor raises the probability of being in the same regime the following period. To determine the average margin before and after the introduction of a price floor, we therefore need to take account of the floor's effect on average margins in each state (e.g. making price wars less intense) as well as on the average fraction of the time the market will spend in that regime (e.g. price wars last longer.) This average will be given by

$$\begin{aligned}&\frac{\alpha_0+\gamma_0\overline{W}}{1-\rho_0}\cdot\frac{1-\Lambda(\varphi_1+\omega_1\overline{W})}{2-\Lambda(\varphi_0+\omega_0\overline{W})-\Lambda(\varphi_1+\omega_1\overline{W})}+\\&\frac{\alpha_1+\gamma_1\overline{W}}{1-\rho_1}\cdot\frac{1-\Lambda(\varphi_0+\omega_0\overline{W})}{2-\Lambda(\varphi_0+\omega_0\overline{W})-\Lambda(\varphi_1+\omega_1\overline{W})}\end{aligned}\tag{3}$$

before introduction of the price floor and by

$$\frac{\alpha_0 + \beta_0 + \gamma_0 \overline{W}}{1-\rho_0} \cdot \frac{1-\Lambda(\varphi_1 + \theta_1 + \omega_1 \overline{W})}{2-\Lambda(\varphi_0 + \theta_0 + \omega_0 \overline{W}) - \Lambda(\varphi_1 + \theta_1 + \omega_1 \overline{W})} +$$
$$\frac{\alpha_1 + \beta_1 + \gamma_1 \overline{W}}{1-\rho_1} \cdot \frac{1-\Lambda(\varphi_1 + \theta_1 + \omega_1 \overline{W})}{2-\Lambda(\varphi_0 + \theta_0 + \omega_0 \overline{W}) - \Lambda(\varphi_1 + \theta_1 + \omega_1 \overline{W})} \tag{4}$$

after, where $\overline{W}$ is the average wholesale price.[22]

## 4. Empirical Results

The results of the estimation of equations (1) and (2) by maximum likelihood are presented in Table 3. The first two columns present estimates under the restriction that $\omega_0 = \omega_1 = 0$; the last two columns present estimates of the unrestricted model. We test the fit of the above models using the score-based tests proposed by Hamilton (1996). Results are shown in Table 4. These tests have power against unmodeled serial correlation and heteroscedasticity in equation (1). They also test for omitted higher-order Markov dependence in equation (2). Since a higher-order Markov chain may be rewritten as a first-order chain with a larger number of states, the latter test also has power against omitted states. We test for each effect in each regime and also present a test for the joint null hypothesis of none of these forms of misspecification in any of the model's equations.

**[Tables 3 and 4 here]**

The joint misspecification test in Table 4 finds some evidence of misspecification only in the case where $\omega_0 = \omega_1 = 0$. However, this evidence appears to be entirely confined to higher-order Markov dependence in the collusive regime; the test statistics are more than three times larger than those for any other single test and are the only ones significant at the 5% level. There is no significant evidence of misspecification in the unrestricted model. Although we have provided no theoretical justification for the inclusion of wholesale prices in the regime switching model, we include these results to check whether our conclusions are robust to minor changes in specification, and whether the limited misspecification detected in our restricted model seriously affects the reliability of our statistical inferences. We also find that the effects of wholesale price changes on margins have intuitive interpretations in the context of price wars.[23]

### *The Effects of Regulation on Price Wars*

As far as the effects of regulation are concerned, results are similar for both specifications shown in Table 3. Significant effects of regulation are found in the margin equation as well as in the state transition probability equation, but only in the price war regime. The introduction of a price floor significantly increases margins during price wars and significantly increased the persistence of price war periods.

As already mentioned, in both specifications and in both sets of equations (margin and regime selection) considered, the parameters associated with the regulation dummy ($R_t$) are positive and significant only during price wars. This result is consistent with the theoretical predictions where a price floor regulation has no effect on prices (or margins) during collusive periods since a price close to the monopoly level is assumed to be charged during those periods. Moreover, our results show that during price war regimes, the price floor regulation increases both margins and probabilities of continuing the price war. Table 5 reports estimated regime-dependent conditional probabilities (using equation 2), expected durations (using $1/(1-q)$, where $q$ is the regime dependent conditional probability) and margins (using equations 3 and 4), all computed from the estimates presented in Table 3.

**[Table 5 here]**

During collusive regimes, the transition probabilities are similar before and after regulation and their estimates are quite robust to the inclusion/exclusion of the wholesale price. However, during price war regimes, transition probabilities are significantly higher after regulation again, regardless of the specification considered. For example, in the unrestricted model the transition probabilities increased from 0.34 to 0.90 after regulation during price wars while the same probabilities stayed at around 0.90 during collusive periods.

Regulation therefore significantly increased the expected duration of price war periods, while reducing sharply the expected duration of collusive periods (although not significantly.) Before regulation, the expected duration of price wars was a little less than two weeks, which rises to ten weeks after regulation.

Results on estimated margins are fully consistent with the theory. On one hand, during collusive regimes, estimated margins are about the same magnitudes with and without a price floor regulation. On the other hand, regulation increases significantly the margins during price wars: from approximately –1.5 cents to 3.5 cents with a price floor. However, and unsurprisingly, we verify that regulation did not raise price war margins to the level of collusive margins. Furthermore, since the price floor raises margins in the low-margin state, but increases the time spent in this state, its net effect on the average margin (i.e. its unconditional expectation) in this market is potentially ambiguous. To answer this question, Table 6 presents estimated unconditional probabilities of the collusive regime (this is 1 – the corresponding probability for the price war regime) as well as the unconditional expected margin. From those figures, it appears that the increase in margins during price wars has been almost exactly offset by the increase in the average duration of a price war, resulting in no significant change in the average margins in the industry. In other words, the price floor regulation had little or no effect on average margins (and therefore prices) even if margins are now higher during price wars, simply because those wars now last longer.

**[Table 6 here]**

## 5. Concluding Remarks

In theory, price wars should last longer under a price floor regulation. This prediction is entirely supported by our empirical findings. The results obtained with a Markov Switching Model using data on the Montreal retail market for gasoline show that the introduction of a price floor regulation reduces the intensity of price wars but raises their expected duration.

Our results support three important conclusions. First, given the coherence between the estimated results and the theoretical predictions, it appears that a model of tacit collusion between firms captures well some important features of the evolution of prices on the retail market for gasoline in Montreal in the period 1994-2001. Second, the government regulation succeeded in reducing the severity of price-wars; we see far less pricing below marginal cost after the introduction of regulation than before. Third, the net effect of the regulation on the competitiveness of the industry is ambiguous. There is no demonstrable change in the average price margin after regulation. This would appear to have benefited gasoline consumers only if one assumes that margins would have risen had regulation not been introduced.

**Table 1 Descriptive Statistics**

| Variable | Mean | Minimum | Maximum | Standard Deviation |
|---|---|---|---|---|
| Price : $P_t$ (¢/litre) | 28.7738 | 15.4000 | 49.7000 | 7.1932 |
| Wholesale price: $W_t$ (¢/litre) | 24.3781 | 13.7250 | 42.8250 | 6.8470 |
| Margin : $M_t$ (¢/litre) | 4.3957 | -5.8750 | 8.8000 | 2.2452 |
| Margin before regulation | 4.3504 | -5.8750 | 8.7200 | 2.6105 |
| Margin after regulation | 4.4229 | -1.2000 | 8.8000 | 1.9993 |
| Regulation : $R_t$ (0,1) | 0.6274 | 0 | 1 | 0.4841 |

**Table 2 Percentage of time with margins below or above mean margin**

| | Before regulation (156 weeks) | After regulation (260 weeks) | Total (416 weeks) |
|---|---|---|---|
| Margin <Mean Margin | 43% | 48% | 46% |
| Margin > Mean Margin | 57% | 52% | 54% |

Figure 1 Retail Margins, Montreal 1994-2001

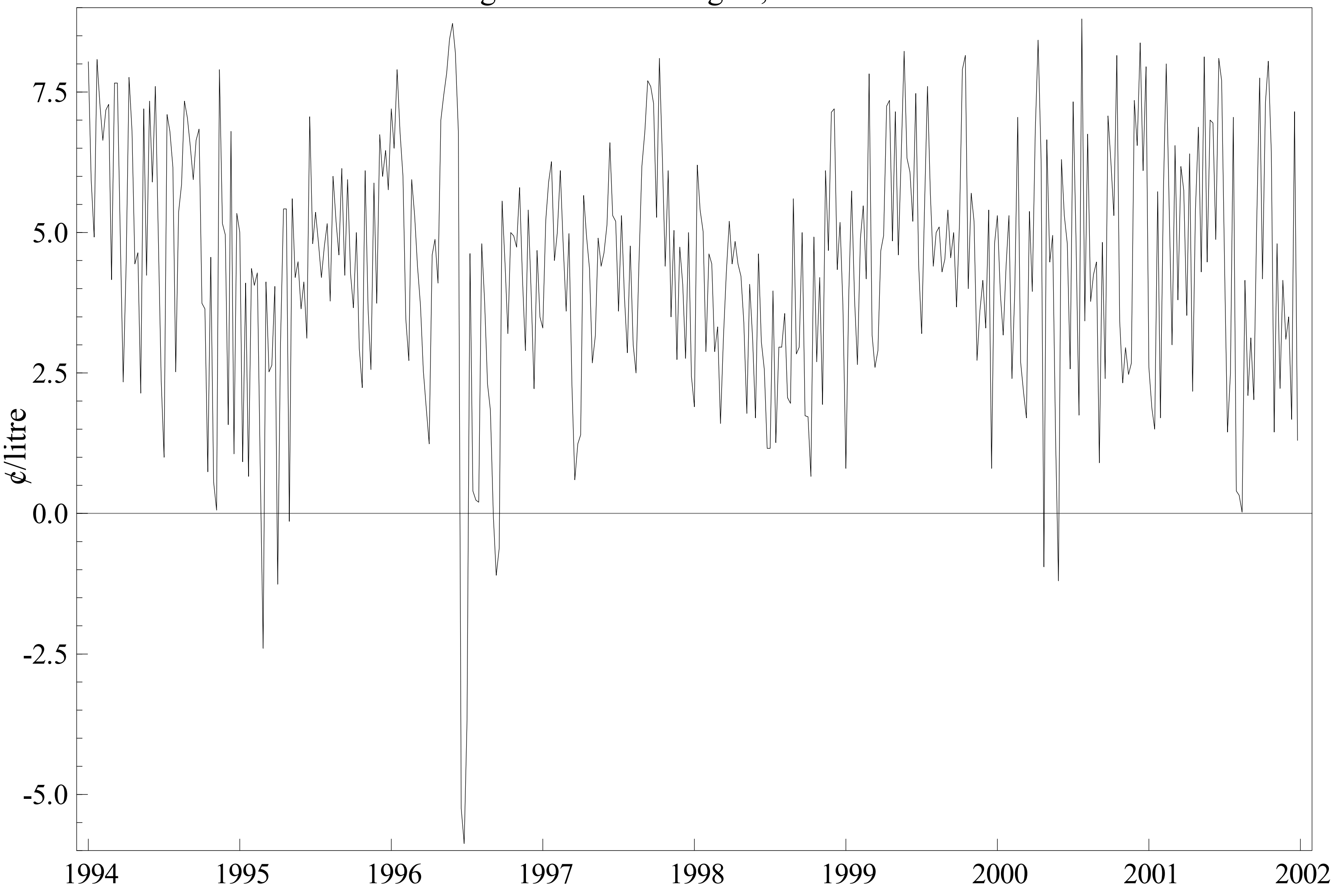

Figure 2 Retail Margins, Montreal 1996-24 - 1996-28

**Table 3 Parameter Estimates (Dependent variable is $M_t$)**

| Parameter | Variable | AR(1) without W Estimated value | AR(1) without W Standard error | AR(1) with W Estimated value | AR(1) with W Standard error |
|---|---|---|---|---|---|
| **Margin Equation** | | | | | |
| $\alpha_{collusion}$ | *constant* | 3.7034* | 0.5926 | 4.0761* | 0.3621 |
| $\rho_{collusion}$ | $M_{t-1}$ | 0.2546* | 0.0947 | 0.2088* | 0.0598 |
| $\beta_{collusion}$ | $R_t$ | 0.1679 | 0.5060 | 0.7352 | 0.4501 |
| $\gamma_{collusion}$ | $W_t$ | | | 0.0067 | 0.0324 |
| $\alpha_{pw}$ | *constant* | -1.3040 | 0.9118 | -1.1864* | 0.5928 |
| $\rho_{pw}$ | $M_{t-1}$ | 0.1242 | 0.1107 | 0.2591* | 0.0739 |
| $\beta_{pw}$ | $R_t$ | 4.4870* | 1.1564 | 3.7059* | 0.6684 |
| $\gamma_{pw}$ | $W_t$ | | | -0.0684* | 0.0214 |
| **Regime Selection Equation** | | | | | |
| $\phi_{collusion}$ | *constant* | 1.6296* | 0.3169 | 1.1344* | 0.2617 |
| $\theta_{collusion}$ | $R_t$ | -0.3362 | 0.6831 | -0.1476 | 0.4610 |
| $\omega_{collusion}$ | $W_t$ | | | -0.1078* | 0.0343 |
| $\phi_{pw}$ | *constant* | -0.0950 | 0.3983 | -0.4228 | 0.3535 |
| $\theta_{pw}$ | $R_t$ | 1.3906* | 0.7181 | 1.7070* | 0.4707 |
| $\omega_{pw}$ | $W_t$ | | | -0.1451* | 0.0511 |
| **Error Variance** | | | | | |
| $\sigma_\varepsilon$ | | 1.8386* | 0.1233 | 1.6395* | 0.0817 |

* : Statistically significant at the 5% confidence level; "*collusion*" stands for collusive regime and "*pw*" stands for price war or competitive regime.

**Table 4 Diagnostic Tests**

| Test for[24] | Without W | | With W | |
|---|---|---|---|---|
| | Statistic | p-value | Statistic | p-value |
| Serial Correlation – Collusion Regime | 0.13 | 0.718 | 3.45 | 0.063 |
| Serial Correlation – Price War Regime | 0.20 | 0.655 | 0.00 | 0.976 |
| ARCH | 5.76 | 0.016 | 2.59 | 0.107 |
| Higher-order Markov Dependence – Collusion Regime | 18.17 | 0.000 | 3.56 | 0.059 |
| Higher-order Markov Dependence – Price War Regime | 3.21 | 0.073 | 0.04 | 0.842 |
| Joint Test | 46.75 | 0.000 | 10.50 | 0.062 |

**Table 5 Regime Dependent Statistics**

| Regime | $P(S_t=i \mid S_{t-1}=i)$ | | E(duration) | | Estimated Margins | |
|---|---|---|---|---|---|---|
| | without W | with W | without W | with W | without W | with W |
| **Collusive** ***($R_t$=0)*** | 0.94841 | 0.87170 | 19.382 | 7.794 | 4.9683 | 5.1520 |
| **Collusive** ***($R_t$=1)*** | 0.90206 | 0.83814 | 10.210 | 6.178 | 5.1936 | 6.0813 |
| **Price War** ***($R_t$=0)*** | 0.46217 | 0.33624 | 1.859 | 1.507 | -1.4889 | -1.6013 |
| **Price War** ***($R_t$=1)*** | 0.90246 | 0.90047 | 10.252 | 10.048 | 3.6344 | 3.4007 |

**Table 6 Unconditional Probabilities and Margins**

| State | $P(S_t$=collusive) | | Estimated Margins | | Sample Margins |
|---|---|---|---|---|---|
| | without W | with W | without W | with W | |
| ***$R_t$=0*** | 0.91247 | 0.83801 | 4.4031 | 4.0581 | 4.3504 |
| ***$R_t$=1*** | 0.49899 | 0.38077 | 4.4124 | 4.4214 | 4.4229 |

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[1] Quoted from the *Report of the Liberal Committee on Gasoline Pricing in Canada*, June 1998 (p. 5).
[2] Translated from the *Mémoire de l'Association Québécoise des Indépendants du Pétrole*, June 1998,
[3] In Rotemberg and Saloner (1986), the punishment price is thus exactly equal to the price floor we consider in this paper, and never appears in equilibrium. Note that all models of collusion which adopt Friedman (1971)'s formulation of "grim" strategies (i.e. punishments of infinite duration) rule out the possibility of punishment prices below marginal costs.
[4] An empirical test by Ellison (1994), in a reexamination of the Joint Executive Committee, gives weak support to Rotember and Saloner (1986)'s theory. Borenstein and Shepard (1996), by using panel data on retail margins on gasoline markets in a sample of U.S. cities, offer evidence in support of Harrington and Haltiwanger (1991)'s theoretical prediction.
[5] In this paper, we follow Slade (1992, footnote 4, p. 260) by using the term "price war" to denote a punishment strategy of some kind in which prices are below their collusive level.
[6] In a second price pattern, retail prices remain "fixed" for months at a time. In a third, less common, and so-called "normal" pattern, retail prices follow closely wholesale prices. In Noel (2003), retail gasoline prices in Montreal are reported to cycle 66.5% of the time, to be fixed 17.9% of the time, and to display a normal behavior in the remaining 15.6%. These figures (in Table 6, p. 35) appear to be intermediate with respect to the cases of Vancouver (43.7%; 30.0%; 27.2%) and Toronto (83.9%; 14.6%; 1.4%). Interestingly, in a related paper which also examines the Montreal market, Eckert (2003) observes that "prices are not easily classified as being rigid over the entire period, or as displaying a price cycle" (p. 157), by using weekly data over the period 1990-1995 that precedes the introduction of a price floor.
[7] In this theoretical context, although the formal model assumes a wholesale price that remains constant over time, one can intuitively evoke an unexpected cost change in order to explain that retail prices may be observed below marginal costs, for some short time period, on real-world markets for gasoline. See Eckert (2002, p. 64).
[8] Following Slade (1989, footnote 3, p. 309), we distinguish oligopoly supergame models of collusion from alternating-move models of price cycles adapted from Maskin and Tirole (1988), on the grounds that stable periods of high prices are absent in the latter model. This distinction is compatible with Eckert (2002)'s remark that cities in which price cycles are observed may be subject to the exercise of less market power than marketplaces in which cycles are not observed.
[9] In Lambson (1994), firm-specific capacities are introduced to make firms dissimilar, with no particular reference to retail gasoline markets. Eckert (2003) and Noel (2003) mention the assumption that gasoline retailers are capacity constrained, but do not formalize it.
[10] In Häckner (1996), the chosen formulation of the differentiation assumption fits well with the fact that, in retail gasoline markets, product varieties are not only differentiated by brand or service, but also (most importantly) by location. This is emphasized in an empirical analysis by Borenstein and Shepard (1996, p. 430), who argue that retailers compete with local rivals only, and thereby justify the use of oligopolistic settings in urban areas where stations are many.
[11] The most severe punishments are defined as "optimal" in the sense that they lead firms to obtain the highest level of sustainable collusive profits. This follows Abreu (1986).
[12] The optimal symmetric punishment must satisfy two conditions. First, the one-period gain from a deviation from the collusive price must be less than (or equal to) what is lost when the punishment price is charged in the next period. Second, the one-period gain from such a deviation must be less than (or equal to) what is lost by prolonging the punishment by one more period. When these two conditions hold with equality, and for a given discount factor δ, they constitute a system of two equations which is solved by the pair ($p^*$, $\underline{p}$). In the absence of restriction on δ, all firms would charge the joint profit-maximizing price as a collusive price, that is $p^* = p^m$. Häckner (1996) demonstrates that there exists a threshold value δ' such that $\delta < \delta'$ implies $p^* < p^m$.
[13] In Häckner (1996), the constant marginal (and average) cost is normalized to zero (i.e. $c = 0$) for simplicity, and the case of positive marginal costs is treated as a generalization (see Häckner (1996), footnote 13, p. 623).

[14] This claim is not limited to the model specifications chosen by Häckner (1996); it also applies (with a change in the formalization of the differentiation assumption) in a recent contribution by Lambertini and Sasaki (2002). Those authors establish that the strictest punishment is optimal because it requires the lowest critical threshold for the discount factor in order to sustain the joint profit-maximizing collusive price. Using this reinterpretation of the definition of optimal penal codes, they focus on collusion at the monopoly level. In that context, when a price floor constraint that binds is introduced, the unchanged monopolistic price can be sustained (i.e., $p^* = p^m$) by a multiple-period price war.

[15] Porter (1983a) and Green and Porter (1984) do not examine gasoline markets. Porter (1983b, 1985) uses data on the Joint Executive Committee, a railroad cartel organized in 1879 to coordinate prices for transport between Chicago and the East Coast. Porter (1985) recognizes that railroads chose price, not quantities, and that the distinction has some implication for the econometric analysis, but reckons that "the basic ideas from the theory still apply" (p. 415). This point is also discussed in Ellison (1994, pp. 38-39).

[16] Slade (1989) models the process of adjustment to new stationary-equilibrium prices. Slade (1992) builds on this using price-war data to estimate different specifications of the inter-temporal reaction function. In Slade (1989, 1992), continuous strategies are distinguished from discontinuous strategies, such as reversion to the one-shot Nash equilibrium in Green and Porter (1984), in that large (small) deviations lead to large (small) punishments. This formulation is justified by Slade (1987), who produces an empirical test of tacit collusion in the Vancouver retail gasoline market. It is found that continuous strategies provide a better model of the price-war dynamics than discontinuous price-reversions to the one-shot Nash equilibrium. In Slade (1992), again with data on retail gasoline activity in a region of Vancouver, reaction functions that are piecewise-linear in the previous-period prices of rivals are given empirical support. Pricing below marginal cost is not discussed.

[17] A process of adjustment is constructed in which prices converge to a new stationary Nash equilibrium for a game with new demand conditions. Following Slade (1989, 1992), here we may keep characterizing firms' reactions as a "punishment", although there is only myopic behaviour and no "crime".

[18] Note that several models produce price wars at equilibrium with more severe punishments than Nash reversion. For example, Abreu, Pearce, and Stachetti (1986) modify the model of Green and Porter (1984) by replacing the continuum of possible output levels with a discrete production set for each firm. Another example is Segerstrom (1988), who introduces a small probability that a player will deviate, and derives conditions under which the deviating firm will find it optimal to "repent" by reducing its own output level for a certain number of subsequent periods. However, some of their specifications, including the assumption that firms are quantity-setting players with undifferentiated products, do not fit some features of retail gasoline markets. Interestingly, Slade (1989, p. 304) notes that in her price-setting supergame model of collusion, it is possible to substitute a stick-and-carrot price structure for the continuous intertemporal reaction function she considers.

[19] In Figure 2, the stepwise pattern of margins emphasizes the fact that each observation in our dataset may account for a pricing policy that remains constant up to seven days.

[20] Slade (1992) addresses this issue using Kalman filtering methods for unobserved component models. However, those are not quite appropriate here since they are designed for unobserved variables that are continuous, not dichotomous. See Hamilton (1994) or Kim and Nelson (2001) for a discussion of the relationship between Kalman filtering approach and the regime switching techniques we use.

[21] See also Filardo (1994) and Filardo and Gordon (1998).

[22] This average should correspond closely to the sample average of margins before or after regulation.

[23] It is also interesting to note in table 3 that the parameters associated with the wholesale price are negative and significant only during the collusive regimes. The reduction in margins in the face of increased costs during collusive periods is consistent with monopolistic pricing behaviour. It is likely that similar behaviour is not observed during price wars because margins cannot be reduced further. Increasing cost during collusive periods also decreased the persistence of these periods.

[24] Each individual test statistic has an asymptotic $\chi^2(1)$ distribution under the null hypothesis of no misspecification, while the joint test is asymptotically distributed as a $\chi^2(5)$ under the null.