The Power to Pass Taxes on - Tax Incidence in the Canadian Retail Gasoline Market*

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Abstract

Since gasoline has a relatively inelastic demand, raising government revenue via gasoline taxes could appear appropriate as it entails a relatively small deadweight loss. However, gasoline retail is generally a highly concentrated market, hence the assumption of perfect competition might be misleading. Theoretically, in oligopolistic markets taxes can be shifted forward less (more) than proportionally to retail prices; a possibility usually denoted by undershifting (overshifting). Generally, whether there is under/overshifting or not depends on unobservable parameters of the demand and cost functions. However, the degree of tax shifting depends on market structure. Therefore, one should consider the interaction between taxes and market structure. The main contribution of this paper is to devise a novel empirical test that allows to identify whether taxes are under- or overshifted. To our knowledge we are the first to explicitly account for an interdependence of tax shifting and market structure. We empirically assess the effect of market structure on tax incidence in oligopolistic markets using data from 10 Canadian cities between 1991 and 1997 finding that taxes in the Canadian gasoline retail market are undershifted.

< Tables and figures at end >

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1 Introduction

Most countries tax gasoline, although to varying degrees. Since gasoline has a relatively inelastic demand, raising government revenue via gasoline taxes could appear appropriate as it entails a relatively small deadweight loss. However, gasoline retail is generally a highly concentrated market leading to prices that are significantly above marginal cost. Hence, tax incidence analysis based on the assumption of perfectly competitive markets might be misleading.

Whereas in long-run equilibrium under perfect competition taxes are passed on fully to consumers (are fully shifted), the situation in oligopolistic markets is different. Taxes can either be shifted forward less or more than proportionally to retail prices. In this paper we devise a novel empirical test based on observables to identify whether a market presents under-, full or overshifting.

Assessing tax incidence in the gasoline market is of particular concern for policy analysis. First, gasoline is an important input. Moreover, in recent years, gasoline taxation has been one of the main tools to encourage rolling CO2 emissions back.[1] Finally, the international turmoil that started in 2008 has considerably increased pressure on governments all around the world in order to reduce budget deficits. From this point of view, gasoline taxes might play an important role as a source of revenues for both local and central governments. [2]

In this paper we aim at empirically assessing the effect of market structure on tax incidence in oligopolistic markets. For this endeavour we use a dataset on the Canadian retail gasoline market comprising monthly observations from 10 Canadian cities over the 1991-1997 period. In Canada gasoline is taxed both at the federal and provincial levels. The market is dominated by a few vertically integrated national (global) players implying a significant level of market concentration. This also holds at the local level possibly due to cost advantages from vertical integration or other barriers to entry. Both variation across time (including several provincial tax changes) and across location help us identify our coefficients of interest.

Our results suggest that the direct effect of taxes is not statistically different from 1 indicating (in line with theory) that, under perfect competition, gasoline taxes in the Canadian retail market are fully shifted to consumers. Similarly, as expected, market concentration raises prices. The direct effect of market structure when no tax is implemented is positive and statistically significant. Interestingly, however, the interaction term is negative and statistically significant, implying that taxes are undershifted.

The rest of the paper is structured as follows. In section 2 we present the existing literature in the field. Section 3 presents the theoretical model behind our empirical analysis. Section 4 gives a brief description of the gasoline industry and taxes in Canada. In section 5 we describe our dataset and the empirical framework. In section 6, we discuss our main results. Finally, in section 7, we provide some concluding remarks.

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[1] See, for example, Palazzi (2011).

[2] As mentioned in the OECD Tax Database, currently around 4% of total tax revenue in Canada comes from environmentally related taxes.

2 Literature Review

In perfectly competitive markets, the incidence of an excise tax depends on the relative elasticities of supply and demand. However, if markets are oligopolistic, a tax could be shifted forward less (more) than proportionally to retail prices. In other words, in imperfectly competitive markets taxes could be undershifted or overshifted. A number of theoretical papers have taken up this point.[3]

The conventional wisdom belief used to be that taxes are not fully passed to consumers because of the potential contraction of the market the producer would face. Krzyzaniak and Musgrave (1963) were among the first to suggest that a tax could be shifted forward more than proportionally to retail prices.

Katz and Rosen (1985) show that the Krzyzaniak-Musgrave result can be rationalized using conventional neoclassic tools. They find that the assumption that an oligopolistic industry acts as if it is competitive or monopolistic can lead to potential mistakes and they also demonstrate the way in which tax incidence depends on market structure.

Seade (1985) questions the role of the perfect competition hypothesis and introduces a cost-side shifter in the equilibrium solution of an oligopolistic market. He shows that, under a certain level of market concentration, a tax hike may indeed increase the producer's equilibrium net price. He concludes that under oligopoly and assuming linear costs, price overshifting turns out to be a likely scenario. Overshifting will occur "if and only if the elasticity of the slope of inverse demand (E) is greater than 1". [4] Note that this is always the case for an isoelastic demand.

Besley (1989) extends Seade's paper by allowing for entry in the market. His main interest is on welfare effects finding that, with entry, taxes in oligopolistic markets can be welfare improving. Regarding the extent of the shifting of taxes, he finds that there is undershifting if the demand function is concave and there is overshifting if it is convex. Finally, he states that overshifting is more likely under free entry.

Delipalla and Keen (1992) compare the incidence effects of excise and ad valorem taxes in two oligopoly models, with and without free entry. They find that ad valorem taxes strictly dominate specific ones in welfare terms. Further, they find that under both the Generalized Cournot and Free Entry Oligopoly model, specific taxes are more likely to be overshifted than ad valorem ones.

Finally, Anderson et al. (2001) extend Delipalla and Keen (1992) by studying the incidence of ad valorem and excise taxes in an oligopolistic industry allowing for product differentiation and price-setting (Bertrand) firms. They analyze the effects of tax incidence on prices and profits in both the short and long runs. They find that the conditions for overshifting arise depending on the curvature of demand. Interestingly, their results are very close to those of a Cournot model with homogeneous demand.

Summarizing these contributions, the pass-through of taxes or the presence of under- or overshifting generally depends on the functional form of the demand curve and the cost function, i.e parameters that are not directly observable to the econometrician. Based on

[3] A more detailed review can be found in Fullerton and Metcalf (2002).

[4] Seade (1985), p.28.

the theoretical literature above we derive a test to detect whether a market displays under- or overshifiting which depends on observables, namely on the interaction term between taxes and market structure.

Tax incidence and market structure have also been studied from an empirical perspective. Harris (1987) and Karp and Perloff (1989) were among the first to empirically study tax incidence in imperfectly competitive markets. Harris (1987) studies the 1983 federal cigarette tax increase in the United States. He finds that the 36% increase in the real price of a pack of cigarettes (adjusted for general inflation) was mainly explained by the augmentation of prices charged by the major U.S. manufacturers and not by the tax hike in the federal excise tax. Karp and Perloff (1989) focus on the Japanese television market using data on a subset of firms. They combine firm specific information with industry aggregates to obtain estimates of market structure, while allowing for cost differences across types of TV sets. The authors then compute the bias on tax incidence from incorrectly assuming a perfectly competitive market. They find that assuming a competitive Japanese television industry underestimates the consumer tax incidence by around 19%.

Besley and Rosen (1999) examine the incidence of sales taxes for several goods. Specifically, they use prices of 12 specific commodities in 155 U.S. cities over the 1982-1990 period and examine the extent to which differences in tax rates and bases are reflected in prices. They find variation in the shifting patterns. For some commodities they can not reject full shifting but, for others, they find overshifting. Finally, they find that the response of prices to a variation in taxes is quite fast, i.e. there is a lag of about one quarter from the tax variation to the effect on prices.

Delipalla and O'Donnell (2001) analyze the tax incidence of ad valorem and specific taxes in the European cigarette market. They use their results as a method for estimating market power and measuring the degree of competition. Basing their analysis on data from the 12 members of the E.U. prior to its expansion in 1995, they find that commodity taxes are not always fully shifted onto consumers. Indeed, after splitting the sample of countries in two different groups, they find that taxes are overshifted for southern European countries but, on the other hand, there is undershifting in the north of Europe where the pressure of the health lobby is stronger. Moreover, their results reject both extremes i.e. perfect competition and collusive behavior. "Firm's behavior in these markets would appear to be no less competitive than the equivalent Cournot and are probably more competitive than this."[5]

More recently, Bergman and Hansen (2010) study the tax shifting of excise taxes on alcoholic and non-alcoholic beverages at the micro-level in Denmark. They focus their analysis on six episodes of tax changes including both tax increases and tax cuts. They find evidence suggesting large differences across brands and different types of beverages but, in general, tax hikes are overshifted whereas tax cuts are undershifted. They interpret their results as evidence of Danish retailers having substantial local market power.

DeCicca et al. (2010) study the shifting of cigarette excise taxes using data on reported prices paid for cigarettes. They focus on the impact of price-search behaviors and find that cigarette excise taxes may be differentially shifted to different group of consumers.

[5]Delipalla and O'Donnell (2001), p.21.

Excise taxes are shifted at lower rates to consumers who undertake more price search behavior and are shifted at higher rates to non-daily, less addicted and light cigarette smokers.

Three papers have studied tax incidence specifically for gasoline. Chouinard and Perloff (2007) study price differences across U.S. states. They find that most of the increase in national gasoline prices over the 90s is explained by a rise in the price of crude oil. Thus, against conventional wisdom, the authors find that other factors such as taxes and market power did not have a significant impact on the price augmentation observed during that decade. On the other hand, differences among states in retail gasoline prices are mainly explained by variations in taxes and market power.

Giacomo et al. (2009), analyze the incidence of specific taxes in fuel markets and exploit the findings to simulate the effects of government interventions aimed at mitigating oil price fluctuations. They use wholesale prices for gasoline and diesel in the Italian fuel industry over the period 1996-2007. They find evidence supporting that "flexible "taxation schemes focused only on excise taxes do not stabilize the fuel price level. Moreover, they mention policies focused on market structure such as the role of the Antitrust Authority as alternative interventions in order to control price fluctuations.

More recently, Marion and Muehlegger (2011) analyze the pass-through rate of federal and state gasoline and diesel taxes to retail prices focusing on the dependence of the shifting on factors constraining the gasoline and diesel supply chains, i.e. within state heterogeneity in gasoline content requirements, refinery capacity utilization, inventory constraints and variation in the demand for untaxed uses of diesel. They find a linear relationship between changes in taxes and variation in prices. Moreover, concerning gasoline taxes, they find full pass-through of both state and federal taxes. In addition, the authors find that the observed pass-through is immediately translated into final prices.

Contrary to our study, none of these papers considers the interplay between the effect of taxes and market structure explicitly. Indeed, some papers implicitly (or explicitly) make the assumption of perfectly competitive markets.

Finally, the closest study to ours is Sen (2001). He investigates the presence of overshifting in cigarette retail prices using data for 10 Canadian provinces from 1982 to 2002. He focuses on whether the presence of overshifting can be attributed to signs of collusion in the industry. The author finds overshifting of provincial taxes but not of federal ones. Moreover, he cannot attribute the overreaction of prices to provincial taxes directly to collusion. We abstract from the collusion hypothesis, but focus instead on the interdependence of market structure and tax incidence which is, once again, absent in Sen's paper.

3 Theoretical model

We use a simple Cournot oligopoly model mostly based on Besley (1989) and Fullerton and Metcalf (2002) where the N firms in the industry choose their level of output taking prices and taxes as given. Firms are symmetric and, for simplicity, we take the number of firms as fixed, but perform comparative statics. Finally, we concentrate on the effect of excise taxes, the main tax instrument in our dataset.[6]

Let the inverse demand function be:

$$p(\sum q_i) \equiv p(Q) \tag{1}$$

where p is the price, q_i is the output of firm i and $Q(\equiv \sum q_i \equiv q_i + Q_{-i})$ denotes the total output of the industry.[7]

The profit function of firm i is as follows:

$$\pi_i(q_i) = p(q_i + Q_{-i})q_i - c(q_i) - \tau q_i \tag{2}$$

where τ is a specific tax on q and $c(q_i)$ is the cost function.

The usual assumptions are imposed in order to assure a stable symmetric Cournot equilibrium:

Assumption 1. $p(Q) : \Re^+ \rightarrow \Re^+$ is twice continuously differentiable and $p'(Q) < 0$ for all Q such that $p(Q) > 0$.

Assumption 2. $c(q) : \Re^+ \rightarrow \Re^+$ is increasing and twice continuously differentiable with c(0)>0 i.e. all firms face a fixed cost.

Since we are focusing on symmetric equilibria, we can omit the subscripts and derive the first and second order conditions of the profits maximizing firm as follows:

$$\text{FOCs: } p'(Q)q + p(Q) - c'(q) - \tau = 0 \tag{3}$$

$$\text{SOCs: } p''(Q)q + 2p'(Q) - c''(q) < 0 \tag{4}$$

We can thus re-express equation 4 as follows:

$$\frac{p'(Q)}{N}(\eta + N + Nk) < 0 \tag{5}$$

where, as it has been defined by Seade (1980), $\eta = \frac{Qp''}{p'}$ is the elasticity of the slope of the inverse demand and $k = 1 - \frac{c''}{p'}$ measures the relative slopes of the demand and marginal cost curves.

Since $p' < 0$, we get the following necessary and sufficient condition for the SOCs to hold:

$$\eta + N + Nk > 0 \tag{6}$$

Now, given symmetric firms:

$$p(Q) \equiv p(q_i + Q_{-i}) \equiv p(Nq), \tag{7}$$

[6]The model can be extended in many directions such as allowing for entry into the industry (See Besley (1989) and Fullerton and Metcalf (2002)), adding an ad valorem tax (Fullerton and Metcalf (2002)) or applying a conjectural variation assumption (Seade (1980)).

[7]Q_{-i} denotes the output of the industry produced by all the other firms but firm i.

thus, plugging equation 7 into the FOCs, we get:

$$p'(Nq)q + p(Nq) - c'(q) = \tau. \tag{8}$$

Differentiating equation 8 with respect to τ and rearranging, we get:

$$\frac{dq}{d\tau} = \frac{1}{p'(\eta + N + k)} \tag{9}$$

and

$$\frac{dQ}{d\tau} = \frac{N}{p'(\eta + N + k)}. \tag{10}$$

Therefore,

$$\frac{dp}{d\tau} = p'\left(\frac{dQ}{d\tau}\right) = \frac{N}{N + (\eta + k)}. \tag{11}$$

Proposition 1. Under perfect competition, (as $N \to \infty$), taxes are fully shifted onto consumers ($\frac{dp}{d\tau} \to 1$). As the market gets more concentrated, the degree of the shifting increases. Indeed, under monopoly (as $N \to 1$), the degree of the shifting is maximized whether we are in presence of under- or overshifting.

Proof: Inspection of equation 11 immediately proofs proposition 1.

Q.E.D.

Even if the proof is straightforward, proposition 1 is an important implication of the model. Equation 11 confirms that under perfect competition, taxes are fully passed through to consumers. Moreover, it shows that the difference of $\left|\frac{dp}{d\tau}\right|$ is the largest for $N = 1$ i.e. under monopoly, the degree of shifting is always the furthest away from one, whether we are in a situation of under- or overshifting.

Proposition 2. If $(\eta + k)$ is positive, $\frac{dp}{d\tau} < 1$ and the market undershifts taxes. On the other hand, if $(\eta + k)$ is negative, $\frac{dp}{d\tau} > 1$ and we are in presence of overshifting.

Proof: A simple inspection of equation 11 immediately proofs proposition 2 and allows us to classify the cases of relevance for our analysis as follows:

$$\frac{dp}{d\tau} = \frac{N}{N + (\eta + k)} \begin{cases} < 1 \Rightarrow & undershifting \\ = 1 \Rightarrow & fullshifting \\ > 1 \Rightarrow & overshifting \end{cases} \tag{12}$$

Q.E.D.

Proposition 3. The sign of $(\eta + k)$ equals the sign of the cross derivative of prices with respect to taxes and market structure.

Proof:

$$\frac{d\left(\frac{dp}{d\tau}\right)}{dN} = \frac{\eta + k}{(\eta + N + k)^2} \tag{13}$$

As $(\eta + N + k)^2 > 0$ for all η, N and k; the sign of equation 13 equals the sign of $(\eta + k)$. **Q.E.D.**

Proposition 3 summarizes the main contribution of our paper. As mentioned in proposition 2, the sign of $(\eta + k)$ tells us whether we are in presence of under- or overshifting. Unfortunately, both η and k depend on the underlying parameters of the demand and cost functions and are, therefore, unobservable and inherently difficult to estimate. However, taking the derivative of equation 11 with respect to N we get equation 13 which implies that the sign of the cross derivative of prices with respect to taxes and the number of firms is also equal to the sign of $(\eta + k)$. Interestingly, the empirical counterpart of this cross derivative is an interaction term of taxes and market structure. Thus, the sign of the coefficient on the interaction term tells us the sign of $(\eta + k)$ and, consequently, whether a particular market under- or overshifts taxes.

4 The gasoline industry and taxes in Canada[8]

The structure of the gasoline market in Canada is composed of three different categories of firms: major vertically-integrated, regional and independent firms. Vertically-integrated firms such as Petro-Canada, Shell and Esso conduct crude exploration, production and development operations as well as downstream refining and retailing operating at a national level. Regional firms also conduct integrated upstream and downstream activities but are geographically limited. For instance, Irving Oil and Ultramar operate in eastern Canada, whereas Husky is located in the Prairies and the West. Together, majors and regionals account for roughly 80% of all retail sales. Finally, independent firms (e.g. Cango in Ottawa and Domo in Vancouver) and super store retailers (Canadian Tire, Real Canadian Superstore, Save on Foods and Costco) do not own refineries and exclusively conduct downstream retailing.

Serious problems seem to exist between refiners and independent retailers. Unlike the U.S., where there are multiple suppliers (many of whom operate only at the wholesale level), Canada has a reduced number of refiners who sell their product to retailers. Moreover, refiners are also involved in retail operations competing with their independent customers. Independents believe that branded dealers are regularly provided with the product at preferential prices (or guaranteed margins) giving them an unfair advantage at the retail level.

[8]This section is based on The Conference Board of Canada, 2001.

Focusing on retailers, they considerably decreased during the 90s from 22.000 in 1989 to 13.250 in 2000, i.e. a reduction of around 40%. Although majors have led the rationalization in terms of retail outlets and independents increased their proportion of retail sites, the market share of the latter ones decreased from 23% in 1990 to 18% in 1999 suggesting an increase in the degree of market concentration.[9]

As stated by the Conference Board of Canada's report, "The gasoline industry in Canada has a limited number of key players who, through their vertical integration and sheer size, are often expected to have power in the marketplace [and] differences in gasoline prices between cities are generally influenced by the different competitive conditions found at the street level."[10]

Finally, as mentioned in the same report, besides the usual barriers to entry (such as economies of scale) that are usually present in the gasoline industry, costs to both entering and exiting the retail business considerably increased during the 90s. In addition, capital requirements for entry were considerable. Moreover, new environmental regulations were implemented increasing the cost of opening a retail outlet and entering the retail business was also harder due to the decrease on retail margins observed throughout most of that decade.

Taxes are the largest component of the pump price in Canada. In 2000, retail gasoline prices consisted of three major components: taxes, crude oil and a refining/marketing component with taxes representing around 42% of the price on average.

Gasoline taxes in Canada can be divided as follows: a federal excise tax, provincial consumption taxes and the Goods and Services Tax (GST) as well as the Provincial Sales Taxes (PST), where applicable. The fiscal structure of the Canadian gasoline market in 2000 is very well summarized by the Conference Board of Canada's report.[11] The federal excise tax is a flat tax imposed across Canada and added on to the other price components of gasoline (crude oil, refining margin and retail margin). In addition, a consumption tax on gasoline is also levied by provincial governments. The provincial consumption tax also is a flat tax, but it varies among provinces. Moreover, The GST is levied on all components of the price of gasoline (including the federal excise tax and the provincial consumption tax). All Canadian provinces pay the GST, although for New Brunswick, Nova Scotia and Newfoundland, it is part of the Harmonized Sales Tax (HST). Quebec is the only province in Canada to explicitly charge a Provincial Sales Tax. The PST is an ad-valorem tax and is calculated on the total cost of gasoline, which includes the ex-tax price of gasoline, the federal excise tax, the provincial consumption tax and the GST. Finally, some Canadian cities apply their own transit taxes which are flat taxes on retail gasoline.

To sum up, although roughly 80% of sales belong to national and regional retailers, concentration varies across cities and time likely due to differences in market shares of

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[9]See also Sen and Townley (2010).

[10]Conference Board of Canada, p.iii.

[11]Even if the situation could have changed during last decade (for instance, in 2007, taxes in Canada represented, on average, 32% of the pump price), the Board's report is really useful for our study because it analyzes the evolution of the gasoline sector in Canada up to year 2000, covering therefore the 1991-1997 period which is the relevant one in this paper.

independent and super store retailers. Moreover, gasoline is taxed both at the federal and provincial levels. Provinces are completely free to set their taxes, implying significant variation across cities. Hence, Canada presents an almost ideal setting to test our theory.

Interestingly, by simply analyzing our raw data, we observe significant cross-city differences in retail price levels. Average prices ranged from 48.86 ¢ in Calgary to roughly 61.40 ¢ in St. John's over the sample period. Regarding market structure, we observe some interesting features in the local retail market across cities and over time. We detect important cross-city differences in market structure, with city Hirschman Herfindahl Index (HHI) averages ranging from 1356 in Ottawa to 2763 in Saint John.[12] Finally, concerning taxes, for our main tax variable (*excise*) we also observe important differences across cities ranging from a 18.01 ¢ average in Calgary to 24.63 ¢ in St. John's.

In Figures 1 and 2 we divide Canada in four regions (Western Canada, Ontario, Quebec and Eastern Canada) and plot the evolution of market structure and taxes for each of them. While the evolution of market concentration seems to be different among regions, excise taxes incremented in all the four regions when comparing the two end points of the sample. Figure 1 shows a general increasing trend of market concentration in Western Canada, Ontario and, particularly, in Quebec. Eastern Canada is the only region where the level of market concentration had decreased during the analyzed period. As shown in figure 2, the evolution is much more homogeneous in terms of excise taxes.

Moreover, we would expect regions with higher Herfindahls to have higher average retail prices as well. Indeed, the correlation coefficient between final prices and market concentration (0.267) is positive and significant at the 5% level. This is confirmed in figure 3. Going back to a city-level analysis; Saint John, St. John's and Halifax with average Herfindahl values of 2763, 2098 and 1974 have the highest average retail prices of 56.63 ¢, 61.40¢ and 56.11 ¢ respectively. Similarly, cities with lower Herfindahls have lower average monthly retail prices. For example, Ottawa and Calgary with average Herfindahl values of 1356 and 1403, show average monthly retail prices of 55.28 ¢ and 48.86 ¢ respectively.

Finally, we expect a similar relationship between retail prices and taxes. The correlation coefficient between retail prices and taxes (0.427) is also positive and significant at the 5% level. Thus, cities with higher taxes should have higher average retail prices and viceversa. St. John's, Quebec and Montreal with the highest federal and state excise taxes (24.63 ¢, 23.58 ¢ and 23.58 ¢ respectively) also have the highest average retail prices (61.40 ¢, 59.79 ¢ and 59.49 ¢ respectively). Similarly, Calgary has the lowest retail price (48.85 ¢) and also is the city with lowest taxes (18.01 ¢).

5 Empirical framework and data

5.1 Data

Our dataset consists of monthly observations for the 1991-1997 period ten of the main cities in Canada: Calgary, Halifax, Montreal, Ottawa, Quebec, Saint John, St. Johns,

[12]As an illustration, the European Commission merger guidelines presume competitive concerns in industries whith an HHI above 2000.

Toronto, Vancouver and Winnipeg.

Our three main variables are those concerning gasoline prices, taxes and the gasoline retail market structure. Regarding retail prices (the dependent variable), MJ Ervin & Associates conducts the Weekly Pump Price Survey (WPPS), a weekly survey of prices for retail gasoline in more than 50 Canadian cities. Though, because of the periodicity of the variation of the other two variables mentioned before, we base our results on monthly averages of gasoline retail prices. Our main tax variable, *excise*, accounts for federal and provincial excise taxes.[13]

In addition to publicly available information such as data on gasoline prices and taxes, we have an interesting dataset containing information on market concentration that allows us to capture the effect of local retail market competition constructed from firm specific sales of regular grade gasoline. These data and the number of stations were obtained from The Kent Group, a private consulting group providing this information from all outlets in the markets it surveys.

We include control variables such as gasoline wholesale prices and the number of gasoline stations in order to control for supply-side characteristics. Wholesale price is also reported in the WPPS and is highly correlated with crude oil price suggesting that it captures part of the effects coming from both the Canadian wholesale and the international crude oil markets. Thus, we decided not to include crude oil price as a regressor. It is worth mentioning that we also add wholesale prices lagged one and two periods in order to control for potential serial correlation i.e. wholesalers might want to adjust their mark ups when retail prices are higher.

Finally, we also control for demand side aspects by including per capita gasoline consumption, average income at the provincial level and the usual demographic controls (provincial unemployment rates, population, percentage of youths between 15 and 24 years old). Table 1 presents summary statistics.

5.2 Empirical framework and econometrics

In order to go beyond simple correlations to estimate the impact of gasoline taxes and market structure on retail prices, we estimate the following model:

$$Price_{it} = \beta_0 + \beta_1 Tax_{it} + \beta_2 HHI_{it} + \beta_3 Tax_{it} * HHI_{it} + \mathbf{X}'_{it}\beta + \epsilon_{it}, \qquad (14)$$

where $Price_{it}$ denotes retail prices including taxes. Tax_{it} accounts for both federal and provincial excise taxes. HHI_{it} is the usual Hirschman Herfindahl Index included to capture market concentration and, finally, $\mathbf{X}_{it}$ is the set of control variables described before.

To the best of our knowledge, the incidence of gasoline taxes has usually been studied based on linear-additive models. As mentioned before, we are interested in the interdependence of the effects of taxes and market structure. We thus base our research on a

[13]As a robustness check, we also compute the estimations using two additional tax variables: *state* that only accounts for excise taxes at the provincial level and *both*, a broader tax variable including federal and excise taxes and the goods and services tax (GST) which was converted into dollar equivalents because it was originally expressed as a percentage of final prices. Results do not significantly differ and are available upon request.

multiplicative interaction model. In general, the presence of an interaction term implies that the impact of one of the variables depends on the level of the other variable included in it. Our main hypothesis is that the shifting of taxes to final prices might depend on the level of market concentration. Thus, in addition to the direct effects of taxes and market concentration on retail prices one should take into account their interdependence. Moreover, when using multiplicative interaction models, the interpretation of the coefficients varies i.e. the coefficients do not indicate the average effect of a certain covariate on the dependent variable as in an additive model. This will play an important role when interpreting our results in the following sections.

To say it differently, previous studies on tax incidence measure the direct effect of taxes through the coefficient β_1, e.g. a coefficient larger than one implying overshifting. Similarly, in these models, β_2 measures the direct effect of market concentration on retail prices. In our case, β_1 measures the direct pass through effect of taxes under perfect competition, i.e. when $HHI = 0$ and β_2 accounts for the direct effect of market structure on prices when no tax is implemented. Finally, β_3, the coefficient of the interaction term (absent in previous studies), measures the effect of market concentration on the degree of tax shifting. We thus expect β_1 to be equal to 1 and β_2 to be positive. Finally, the sign of β_3 - that indicates whether we are in presence of under- or overshifting - is theoretically ambiguous and hence, its sign remains an empirical issue.

In Figure 4, we plot the different prices and the corresponding taxes for the less and more concentrated cities i.e. cities showing a level of concentration under or above the median HHI. Interestingly, we observe a difference in the slope of both fitted values lines; less concentrated markets having a steeper fitted values line than more concentrated ones. Thus, a first look at our data, suggests that $\beta_3 < 0$ and, therefore, that taxes are undershifted.

5.2.1 Serial Correlation, Clustering and Bootstrapping

Our basic specification presents robust standard errors. However, it could be warranted to allow for other standard errors structures. A typical issue when working with time series (or panel data) is the presence of serial correlation. By producing Newey-West standard errors we assume the error structure to be heteroskedastic and possibly autocorrelated up to some lag. Other strategy in the same direction is to fit our panel-data linear model by using feasible generalized least squares. This strategy allows us to estimate our model in the presence of AR(1) autocorrelation within panels and cross-sectional correlation and heteroskedasticity across panels.

Finally, another possibility is to allow the error terms to be correlated within clusters. Clustering at the city level and computing cluster-robust standard errors seems straight forward in our model. Indeed, as it has been shown by Moulton (1990) and is well known from the literature in the field,[14] the estimation of the covariance matrix without controlling for clustering can lead to understated standard errors and overstated statistical significance. Under the usual assumptions, the OLS estimator is unbiased in small samples

[14]See also Moulton (1986), Wooldridge (2003) and Stock and Watson (2008) among others.

and normally distributed or consistent and approximately normally distributed when we work with a large number of clusters. However, when the number of clusters is not big enough, the OLS estimator is not efficient any more and the usual standard errors of the OLS estimator and tests based on them are not valid. In other words, the cluster-robust covariance matrix is consistent when the number of clusters tends to infinity. Although in practice a minimum of 50 clusters seems to be enough to ensure consistency, we are working with only 10 clusters. Thus the standard procedures for clustering might not be valid. Moreover, with a low number of clusters, bootstrap re-sampling methods produce a finite number of possible pseudo-samples. Thus, the bootstrap distribution will not have the desired smoothness even when replicating many times. In addition, for some pseudo-samples either the coefficient or the standard errors may be inestimable.

An alternative in order to solve this issue is the application of the wild bootstrap method applied in Cameron et al. (2008). The authors find that the wild cluster bootstrap performs really well even with six clusters and without a noticeable loss of power after accounting for size.

5.2.2 Simultaneous Causality Bias

Finally, we detect two potential simultaneous causality bias issues to be addressed. First, higher prices could encourage entry and thus impact market structure. Therefore, market concentration may be endogenous to prices. Second, coefficient estimates of the effects of taxes may be confounded because of simultaneity bias between taxes and prices. In other words, tax setting authorities might adjust tax rates given current prices.

We apply Instrumental Variables estimation in order to solve both endogeneity issues. First, concerning market structure, we can expect P_{it} and HHI_{it} to be correlated but we can be fairly sure that $Price_{it}$ and HHI_{it-1} or HHI_{it-2} are not. Thus, we use the one- and two-periods lagged value of HHI_{it} as an instrument to correct this potential bias.

Similarly, we want to control for the possible endogeneity of taxes mentioned above. In this case, we follow the strategy implemented by Besley and Case (2000) and construct instruments based on the political party in power at the state level. As instruments, we use data indicating whether the political party in power is the same at both the federal and state levels and the share of seats of the political party in power in each province. The idea behind our strategy is that political parties may influence the set of policies to be implemented (including taxation) without having any direct causal relationship with retail prices.

6 Estimation and results

6.1 Main results

Table 3 shows our main results.[15] The direct effect of excise taxes on prices is, under perfect competition, statistically significant and positive in all the specifications. Similarly, the direct effect of market structure on prices (without taxes) is, as expected, statistically significant and positive as well. Finally, our the coefficient of the interaction term indicates that taxes are undershifted.

Interestingly, our results are robust to the different specifications we explore. By applying a simple panel data - fixed effects approach,[16] (Column 1) suggests that, when $HHI = 0$, taxes are overshifted to final prices. $\beta_1 = 1.361$ is statistically significant and bigger than 1, thus, under perfect competition, a 1$ augmentation of excise taxes results in the final price increasing by 1.361$. This result seems to contradict theory that suggests a full shifting of taxes under perfectly competitive markets. Nevertheless, when computing a test for fully shifting, we cannot reject the null hypothesis that $\beta_1 = 1$ i.e. we cannot reject full shifting[17]. Interestingly, our results do not vary considerably when controlling for other factors such as seasonality or when applying a time trend (columns 2 and 3 respectively). In fact, in both columns, β_1 slightly decreases and we still cannot reject that it statistically equals 1. Finally, in column 4, after controlling for city fixed effects and seasonality we apply a city-specific time trend instead of the general one applied in columns 2 and 3. Again, our results are robust with $\beta_1 = 1.311$.

As expected, β_2 is positive and statistically significant in all the specifications. In line with theory, this suggests that market concentration (when no taxes are implemented) has a positive effect on market prices.

Concerning the coefficient of the interaction term, β_3, it is negative and statistically significant at the 1% level in all the four specifications of our model suggesting that, in our dataset, taxes are undershifted. This is our main result.

In addition, most of the controls show the expected signs and are statistically significant (except for per capita gasoline consumption, which is not significant in none of the models).

As mentioned in section 5, previous studies on tax incidence are based on linear-additive models. Table 2 shows how our results would have been based on such a model. The effect of taxes on prices is (in all the four specifications) statistically significant and smaller than 1 implying a certain level of undershifting. This is in line with our results. Similarly, all the control variables show the expected sign and are statistically significant;

[15]In order to simplify the interpretation of the estimates, we re-scale the variable variable HHI. Hence, in all the four tables, the range of HHI is defined between 0 to 1 and not between 0 and 10.000 as it has been indicated in previous sections.

[16]Following Hausman (1978) and Baltagi (2009) we compare the $\hat{\beta}_{GLS}$ and $\hat{\beta}_{Within}$ estimators in order to choose among the Fixed Effects and Random Effects models. We perform a Hausman's Specification Test rejecting the null hypothesis that the RE is consistent and, thus, implement the FE approach as our main specification. RE is applied as a robustness check.

[17]β_1 is close to but different from 1. This could possibly be the case because we are estimating out of sample i.e. our market is not competitive, thus $HHI = 0$ is not a relevant case in our sample.

except for per capita gasoline consumption, which is not significant in none of the models (exactly as in 3). Finally, the effect of market structure shows the opposite sign but it is not significant in neither of the specifications.

Finally, as suggested in Brambor et al. (2006), we produce a figure that helps us giving an interpretation to our main empirical result, i.e. the interactio term's coefficient. In figure 5 we plot the interdependence of tax incidence and market structure i.e. we show how the level of tax incidence varies with the level of market concentration.[18] The sign of the coefficient of the interaction term, β_3, represents the slope of the line indicating whether taxes are over- or under shifted. When both, the upper and lower bounds of the confidence interval are under (above) 1, we are in presence of undershifting (overshifting). When this is not the case, we can not reject full shifting. As it can be observed in the figure, under prefect competition we can not reject full shifting. This is in line with theory. Similarly, as theory suggests, the degree of undershifting increases with market concentration as we get closer to the monopoly scenario i.e. as $HHI_{it} \rightarrow 1$. In other words, taxes in the Canadian gasoline retail market are undershifted to consumers, the degree of the shifting increasing with market concentration and, under perfect competition, we cannot reject full shifting.

6.2 Robustness checks

In tables 4 and 5 we show some of the robustness checks we compute. The former mainly aims at checking the robustness of our results by applying different standard errors' structures. Column 1 only replicates column 3 in table 3 which is our benchmark specification. Column 2 applies a RE to our man specification. In columns 3 and 4 we tackle possible autocorrelation issues. As it is frequent when working with time series (or panel data), the series could be correlated with own lagged values. If this is the case, we are in presence of serial correlation (or autocorrelation) and, thus, the OLS estimates will not be efficient anymore. Even when the Wooldridge test for autocorrelation indicates that this should not be an issue in our case, we compute a FGLS estimate as an additional robustness check. In column 3, we apply a regression allowing for Newey-West standard errors. Thus, the error structure is assumed to be heteroskedastic and possibly autocorrelated up to 2 lags. In column 4, we compute a feasible generalized least squares (FGLS) estimator. As it can be observed, the coefficients of the three main variables (*excise*, HHI and the interaction term) have the expected signs and are statistically significant.

In columns 5 and 6 we allow the errors to be correlated within cities by clustering at the city level. Because of the structure of our data, this strategy seems quite straight forward. The coefficients of the three main variables (*excise*, HHI and the interaction term) have the expected signs in both specifications Fixed and Random Effects (columns 5 and 6 respectively). Though, except for Excise in the RE approach, the estimates are not significant any more. Moreover, as mentioned in section 5.3.1, this strategy could present lead to inefficient estimates because of the low number of clusters we are working with. Therefore, we apply the wild bootstrap procedure described in Cameron et al. (2008).

[18]Figure 5 is based on our main specification, i.e. column 3 in Table 3.

Though, as we could expect, applying a more restrictive procedure reduces (even more) the level of significance of our estimates. Our coefficients are still robust in terms of signs but, none of them, is statistically significant. [19]

5 aims at tackling the possible simultaneous causality bias issues described in section 5.2.2. The first three columns are based on FE-2SLS whereas columns 4, 5 and 6 replicate the same specifications under a RE-2SLS approach. In columns 1 and 4, we only tackle endogeneity issues concerning taxes. We instrument the variable our tax variable, *excise*, using all exogenous regressors plus the share of seats of the political party in power in each province, a dummy indicating whether the political party in power is the same at both the federal and state levels. Similarly, in columns 2 and 5 we only instrument the market structure variable, HHI_t, using all exogenous regressors plus HHI lagged 1 and 2 periods (HHI_{t-1} nad HHI_{t-2}). Finally, in columns 3 and 6, we tackle both endogeneity issues at the same time instrumenting both excise and HHI using the the sum of the instruments described before.

Our set of instruments is a valid one. Both the exogeneity and relevance conditions are fulfilled. First, we cannot reject the null hypothesis of the Sargan test of overidentifying restrictions i.e the instruments are uncorrelated with the error term. Second, concerning the relevance of the instruments, we compute a Weak Identification Test following Stock and Yogo (2002)[20]. For columns 1, 2, 4 and 5; where only one endogenous variable is considered at the time, we provide the 1st-stage F-statistic. In all the four specifications the F-statistic is bigger than 10 implying that we are not in presence of weak instruments. In columns 3 and 6, where we consider two endogenous variables silmutaneously, we provide the Cragg-Donald Wald F-statistic to be compared with the values given by Stock and Yogo (2002). In column 3, we reject the null that the set of instruments is weak at a 20% maximal IV relative bias. [21] In column 6, we can reject the null that the set of instruments is weak at a 10% maximal IV relative bias.

Finally, concerning the estimates, all the three main variables (*excise*, HHI and the interaction term) show the expected signs and are statistically significant at the 1% level.[22] Nevertheless, as it is often the case, the FE-2SLS estimates are bigger than the FE-OLS ones. Even if the signs of the estimates are the expected ones and they are all statistically significant, the estimates increase considerably in magnitude and even esxplode in some

[19]Using the wild-cluster bootstrap-t procedure to improve inference while working with few clusters, we find the following p-values: Under the FE approach; 0.25375, 0.26873 and 0.34466 for *excise*, HHI and the interaction term respectively. Under the RE approach the values are the following ones: 0.32068, 0.44555 and 0.69630 for *excise*, HHI and the interaction term respectively.

[20]See also Hausman et al. (2005) and Baum et al. (2007).

[21]It is worthmentioning that, although we do not have other values than the one provided by Stock and Yogo (2002), we would probably also reject the null at a 15% maximal IV relative bias given that the Cragg-Donald Wald F-statistic of 7.93 has to be compared with the following values: 10% maximal IV relative bias, 8.78 and 20% maximal IV relative bias, 5.91.

[22]Following the procedure in Baltagi (2009), we perform an alternative Hausman test based on the contrast between the fixed effects 2SLS and the random effect 2SLS estimators. The idea is that, after controling for endogeneity issues, the result of our previous Hausman test could have changed. Nevertheless, once again, we reject the null that the RE estimator is consistent. Hence, the FE specification is the preferred specification. Nevertheless, we show both the FE-2SLS and RE-2SLS estimations as an additional robustness check.

cases such as columns 5 and 6. Therefore, the interpretation of our results, in particular the FE-2SLS ones, should be done with some caution. This is the reason why our main results are based on the FE-OLS specification and we comopute FE-2SLS only as a robustness check.

7 Concluding remarks

The Canadian gasoline retail is a concentrated market. Thus, the perfect competition assumption might be misleading in order to analyze the effects of taxes on prices. As it has been already shown, in oligopolistic markets, taxes can be shifted forward less (more) than proportionally to retail prices; a possibility usually denoted by undershifting (overshifting). Generally, whether there is under/overshifting or not depends on unobservable parameters of the demand and cost functions. However, the degree of tax shifting depends on market structure. Hence, one should consider the interaction between taxes and market structure.

Following the models used by Besley (1989) and Fullerton and Metcalf (2002) we show that, under perfect competition, taxes are fully shifted to consumers. In addition, the degree of the shifting (whether it is an under- or overreaction to a variation in taxes) is maximized under monopoly. Indeed, our model shows that, whether we are in presence of under-, full- or overshifting depends on the sign of $(\eta + k)$, defined by underlying parameters of the demand and cost functions. Moreover, the sign of $(\eta + k)$ equals the sign of the cross derivative of prices with respect to taxes and market structure. This is in line with the studies mentioned before and with other authors that had studied the effects of tax incidence under oligopolistic markets. Interestingly, in our procedure, the empirical counterpart of the cross derivative of prices with respect to taxes and market structure is the coefficient of the interaction term which is observable to the econometrician. In other words, the sign of β_3 equals the sign of $(\eta + k)$ and, therefore, is an indicator of the presence of under/overshifting in a given market.

Furthermore, one could consider which are the welfare implications of levying a tax in such a simple model. Following Besley (1989),[23] one could consider that the impact on welfare depends both on the changes in aggregate output and in output per firm. Indeed, when evaluating the impact of welfare at the zero tax position i.e. $\frac{dW}{d\tau}\Big|_{\tau=0}$, the impact on welfare only depends on the change in output per firm. Thus, based on proposition 1 in Besley (1989), [24] we can summarize the results in terms of welfare implications as follows: Under free entry and assuming linear costs, introducing a small specific tax on output i) rises prices less (more) than proportionally and ii) increases (reduces) welfare if $(\eta + k) > (<)0$.

Finally, using a dataset on the Canadian retail gasoline market comprising monthly observations from 10 Canadian cities over the period 1991 to 1997, we empirically test whether taxes are under or overshifted. To sum our results up, we find that taxes are

[23]For a complete analysis on welfare implication, see also Seade (1985).

[24]Besley (1989), p. 366.

undershifted to consumers, that the degree of the shifting increases with market concentration and that, under perfect competition, we cannot reject full shifting.

Our main contribution is to devise a simple test of tax under/overshifting. The simple framework described in this paper is therefore useful in order to detect whether taxes are under or overshifted in a given market. Moreover, because of the simplicity of the data required to apply our procedure, we think that this methodology could be used for academic purposes but also by policy makers in order to asses issues such as those related to taxation policies or collusion. Finally, we see this paper as a first step to asses the impact of market structure on tax competition both at the horizontal and vertical levels. We honestly think that our results open a door to develop future research based on them. In a second step, we aim at linking the vast literature on tax competition[25] with the tax incidence results described in this paper. Another interesting issue to explore is whether and how our results are affected by the presence of different size firms.[26]

[25]See, for example, Hayashi and Boadway (2001), Devereux et al. (2007) and Brulhart and Jametti (2006) among others.

[26]See Vigneault and Wen (2002).

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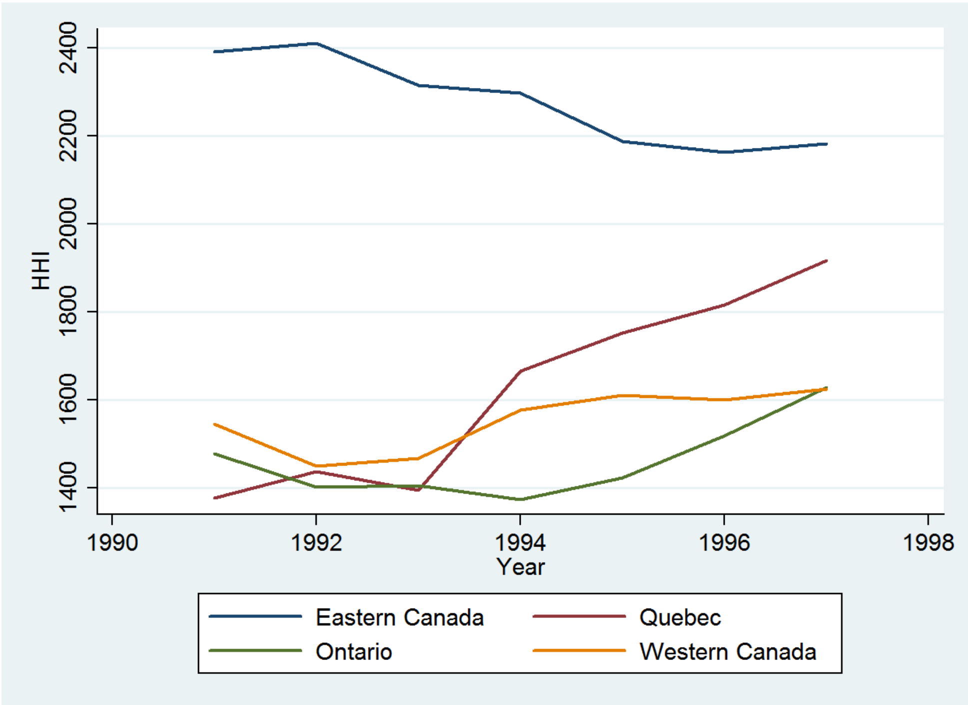

Figure 1: Time Evolution of Market Concentration

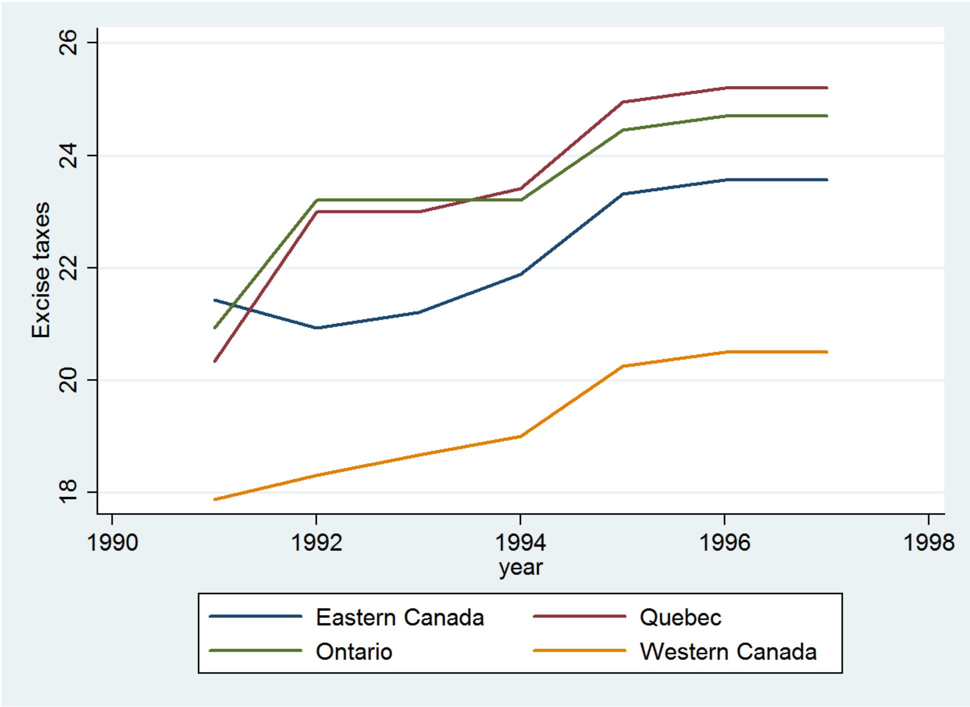

Figure 2: Time Evolution of Excise Taxes

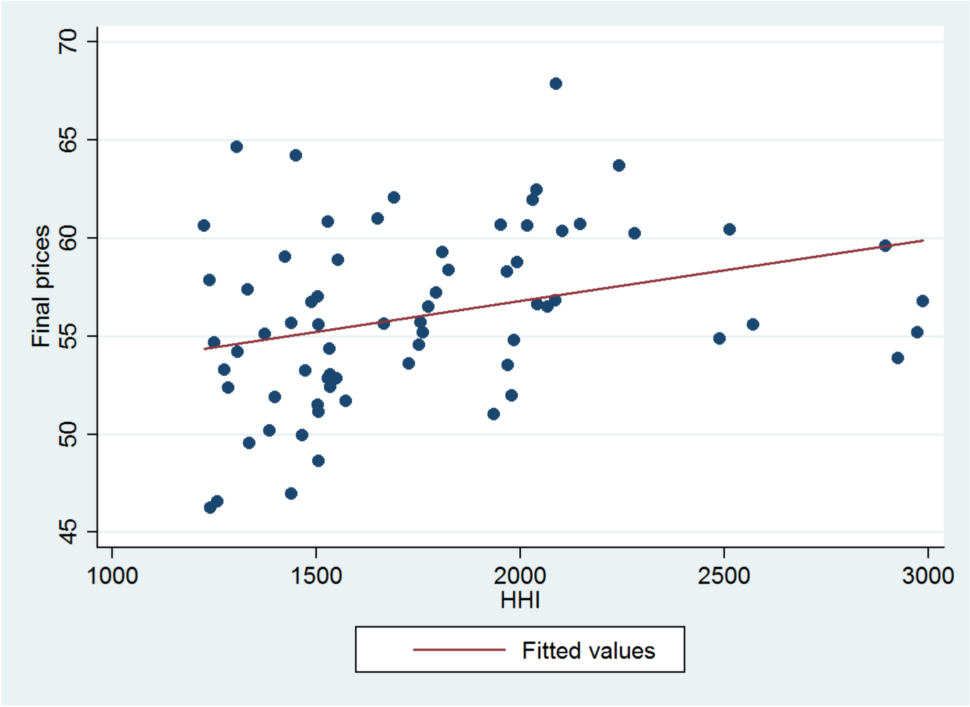

Figure 3: Prices and Market Structure

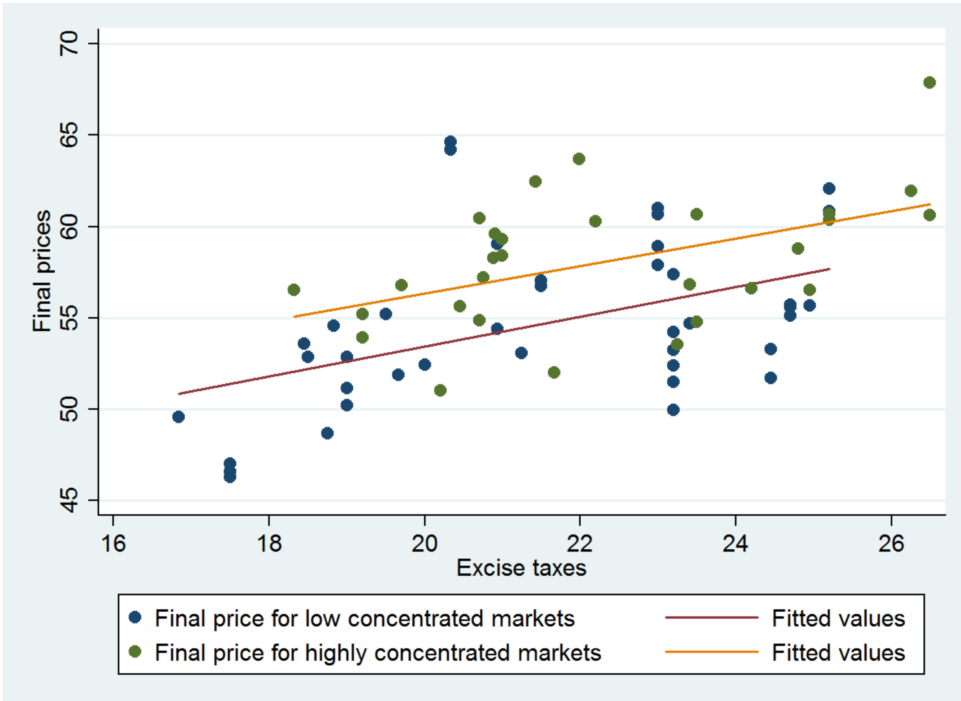

Figure 4: Prices and Taxes for Low and Highly Concentrated Markets

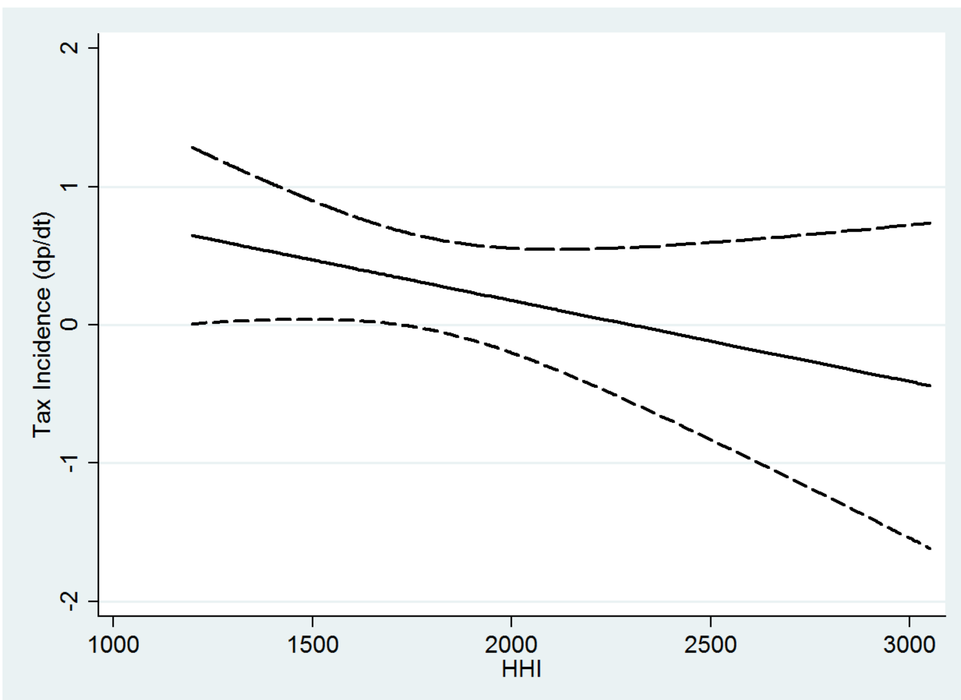

Figure 5: Interdependence of Tax Incidence and Market Structure

Table 1: Summary statistics

Variable	Units	N	Mean	Std. Dev.	Min.	Max.
Final retail price including taxes	¢/litre	840	56.05	5.13	38.67	72.93
State excise tax	¢/litre	840	12.78	2.30	7	16.5
Federal and State excise taxes	¢/litre	840	21.89	2.52	15.5	26.5
Herfindahl Hirschiman Index	0-10.000	840	1766.7	440.5	1196.1	3051.4
Wholesale price of gasoline	¢/litre	840	23.29	2.88	15.24	37.9
Per capita consumption of gasoline	litres (in 1.000)	840	0.20	0.07	0.07	0.50
Total number of stations	1.000	840	0.34	0.35	0.05	1.42
Population	100.000	840	13.22	13.37	1.28	44.99
Prop. population aged 15-24	%	840	14.00	1.74	9.30	19.27
Unemployment rate	%	840	11.39	3.28	5.4	21.5
Average annual income	C$ (in 1.000)	840	44.71	6.69	33.8	55.8

Table 2: OLS and Fixed Effects estimates - Cross city and time-series data (1991-1997)

	(1)	(2)	(3)	(4)
Excise	0.180*	0.188**	0.256**	0.182*
	[0.0954]	[0.0955]	[0.124]	[0.0957]
HHI	-5.199	-5.205	-5.798	-5.107
	[7.802]	[7.745]	[7.661]	[7.743]
Wholesale	0.669***	0.640***	0.644***	0.639***
	[0.0734]	[0.0785]	[0.0791]	[0.0785]
Wholesale_lag1	0.562***	0.540***	0.541***	0.540***
	[0.111]	[0.113]	[0.113]	[0.113]
Wholesale_lag2	-0.241***	-0.225***	-0.229***	-0.225***
	[0.0847]	[0.0833]	[0.0833]	[0.0833]
Unemployment rate	-0.311***	-0.326***	-0.356***	-0.324***
	[0.116]	[0.114]	[0.120]	[0.114]
Population	0.488**	0.542***	0.571***	0.555***
	[0.192]	[0.193]	[0.192]	[0.195]
Prop. aged 15 to 24	-0.221*	-0.218*	-0.264*	-0.214*
	[0.131]	[0.130]	[0.136]	[0.130]
Stations	9.856***	10.54***	9.698***	10.40***
	[3.170]	[3.250]	[3.413]	[3.192]
Gasoline cons. pp	-0.602	-1.785	-1.493	-1.866
	[3.062]	[3.009]	[3.034]	[3.009]
average income	0.226**	0.235***	0.236***	0.236***
	[0.0888]	[0.0866]	[0.0866]	[0.0866]
City FE	YES	YES	YES	NO
Quarter FE	NO	YES	YES	YES
Time trend	NO	NO	YES	NO
City Specific Time trend	NO	NO	NO	YES
Observations	820	820	820	820
Adjusted R^2	0.7433	0.7468	0.7471	0.7467

Intercept included in all regressions.
Standard errors based on robust covariance matrices in brackets.
* p<0.10, ** p<0.05, *** p<0.01

Table 3: OLS and Fixed Effects estimates - Cross city and time-series data (1991-1997)

	(1)	(2)	(3)	(4)
Excise	1.361***	1.304***	1.348***	1.311***
	[0.343]	[0.340]	[0.357]	[0.340]
HHI	127.6***	120.4***	118.5***	121.9***
	[39.18]	[38.96]	[38.69]	[38.93]
Excise_HHI	-6.264***	-5.924***	-5.859***	-5.993***
	[1.793]	[1.793]	[1.782]	[1.793]
Wholesale	0.691***	0.663***	0.666***	0.663***
	[0.0732]	[0.0784]	[0.0790]	[0.0783]
Wholesale_lag1	0.547***	0.527***	0.527***	0.527***
	[0.111]	[0.113]	[0.113]	[0.113]
Wholesale_lag2	-0.204**	-0.191**	-0.194**	-0.190**
	[0.0875]	[0.0858]	[0.0856]	[0.0858]
Unemployment rate	-0.241**	-0.259**	-0.284**	-0.256**
	[0.120]	[0.118]	[0.123]	[0.118]
Population	0.462**	0.516***	0.540***	0.529***
	[0.198]	[0.199]	[0.197]	[0.201]
Prop. aged 15 to 24	-0.271**	-0.265**	-0.302**	-0.261**
	[0.134]	[0.133]	[0.139]	[0.133]
Stations	11.17***	11.76***	11.05***	11.63***
	[3.270]	[3.324]	[3.463]	[3.264]
Gasoline cons. pp	-1.028	-2.124	-1.880	-2.200
	[3.012]	[2.969]	[2.992]	[2.968]
Average income	0.274***	0.281***	0.281***	0.282***
	[0.0872]	[0.0856]	[0.0855]	[0.0856]
City FE	YES	YES	YES	NO
Quarter FE	NO	YES	YES	YES
Time trend	NO	NO	YES	NO
City Specific Time trend	NO	NO	NO	YES
Observations	820	820	820	820
Adjusted R^2	0.7474	0.7505	0.7507	0.7505
Test full shifting	1.10	0.80	0.95	0.83
	0.2933	0.3717	0.3297	0.3613

Intercept included in all regressions.
Standard errors based on robust covariance matrices in brackets.
Test for fully shifting under perfect competition (excise = 1): χ^2 statistic above and P values below.
* p<0.10, ** p<0.05, *** p<0.01.

Table 4: Robustness checks - Cross city and time-series data (1991-1997)

	(1)	(2)	(3)	(4)	(5)	(6)
Excise	1.348***	1.187***	1.209**	1.469***	1.348	1.187*
	[0.357]	[0.218]	[0.475]	[0.324]	[0.847]	[0.676]
HHI	118.5***	93.04***	106.2**	131.5***	118.5	93.04
	[38.69]	[24.37]	[53.81]	[37.37]	[83.40]	[72.66]
Excise_HHI	-5.859***	-3.871***	-5.224**	-6.422***	-5.859	-3.871
	[1.782]	[1.229]	[2.471]	[1.702]	[4.608]	[3.934]
Wholesale	0.666***	0.713***	0.964***	0.764***	0.666***	0.713***
	[0.0790]	[0.0857]	[0.0514]	[0.0701]	[0.0860]	[0.0926]
Wholesale_lag1	0.527***	0.489***	n.a.	0.427***	0.527***	0.489***
	[0.113]	[0.119]		[0.0914]	[0.0742]	[0.0731]
Wholesale_lag1	-0.194**	-0.134	n.a.	-0.170***	-0.194	-0.134
	[0.0856]	[0.0864]		[0.0609]	[0.108]	[0.0962]
Unemployment rate	-0.284**	0.146**	-0.288*	-0.160	-0.284	0.146
	[0.123]	[0.0631]	[0.162]	[0.111]	[0.272]	[0.0892]
Population	0.540***	-0.174***	0.663***	0.405**	0.540	-0.174***
	[0.197]	[0.0200]	[0.255]	[0.168]	[0.462]	[0.0445]
Prop. aged 15 to 24	-0.302**	-0.171*	-0.278	-0.336**	-0.302	-0.171
	[0.139]	[0.102]	[0.196]	[0.136]	[0.291]	[0.172]
Stations	11.05***	7.630***	12.65**	10.72***	11.05	7.630***
	[3.463]	[0.780]	[4.934]	[3.358]	[6.468]	[2.126]
Gasoline cons. pp	-1.880	-22.77***	-1.252	-2.939	-1.880	-22.77***
	[2.992]	[2.014]	[2.848]	[2.881]	[2.279]	[4.846]
Average income	0.281***	-0.130***	0.296***	0.295***	0.281	-0.130
	[0.0855]	[0.0392]	[0.114]	[0.0903]	[0.172]	[0.104]
Cluster (City)	NO	NO	NO	NO	YES	YES
City FE	YES	NO	YES	YES	YES	NO
Quarter FE	YES	YES	YES	YES	YES	YES
Time trend	YES	YES	YES	YES	YES	YES
Observations	820	820	840	820	820	820
Adjusted R^2	0.7507	0.7025		n.a.	0.497	
Test full shifting	0.95	0.74	0.19	2.10	0.17	0.08
	0.3297	0.3904	0.6594	0.1474	0.6906	0.7822

Intercept included in all regressions.
In column 3, the lagged values of Wholesale (Wholesale_lag1 and Wholesale_lag2) are excluded because the Newey-West specification has already been computed considering 2 lags in the autocorrelation structure.
Standard errors in brackets. Columns 1 and 2, based on robust covariance matrices. Column 3, based on Newey-West standard errors. Column 4, based on an heteroskedastic but uncorrelated error structure. Columns 5 and 6, based on cluster-robust covariance matrices.
Column 4 uses a heteroskedastic but uncorrelated error structure. R^2 does not apply.
Test for fully shifting under perfect competition (excise = 1): χ^2 (for columns 1, 2, 4 and 6) and F-statistics (for columns 3 and 5) above and P values below.
* $p<0.10$, ** $p<0.05$, *** $p<0.01$

Table 5: Robustness checks (IV-2SLS) - Cross city and time-series data (1991-1997)

	(1)	(2)	(3)	(4)	(5)	(6)
Excise	1.797***	8.830***	9.088***	1.319***	3.249***	3.345***
	[0.589]	[1.893]	[1.960]	[0.316]	[1.013]	[1.016]
HHI	171.7**	867.6***	897.7***	108.5***	314.2***	324.5***
	[66.28]	[195.6]	[204.2]	[36.18]	[108.9]	[109.2]
Excise_HHI	-8.281***	-41.32***	-42.70***	-4.635**	-14.78***	-15.28***
	[3.018]	[9.119]	[9.499]	[1.805]	[5.376]	[5.392]
Wholesale	0.676***	0.792***	0.797***	0.717***	0.755***	0.757***
	[0.0779]	[0.107]	[0.108]	[0.0852]	[0.0931]	[0.0935]
Wholesale_lag1	0.522***	0.419***	0.416***	0.487***	0.469***	0.468***
	[0.111]	[0.145]	[0.147]	[0.118]	[0.127]	[0.127]
Wholesale_lag2	-0.181***	0.0510	0.0588	-0.130*	-0.0758	-0.0730
	[0.0855]	[0.122]	[0.124]	[0.0858]	[0.0946]	[0.0947]
Unemployment rate	-0.250**	0.166	0.185	0.149**	0.102	0.0993
	[0.127]	[0.169]	[0.173]	[0.0627]	[0.0677]	[0.0689]
Population	0.528**	0.433*	0.426	-0.176***	-0.219***	-0.222***
	[0.197]	[0.283]	[0.287]	[0.0201]	[0.0292]	[0.0293]
Prop. aged 15 to 24	-0.319**	-0.573***	-0.583***	-0.143	0.302	0.325
	[0.139]	[0.202]	[0.207]	[0.111]	[0.252]	[0.253]
Stations	11.89***	21.06***	21.52***	7.643***	7.989***	8.007***
	[3.565]	[6.407]	[6.552]	[0.772]	[0.805]	[0.808]
Gasoline cons. pp	-1.984	-4.437	-4.493	-22.87***	-22.20***	-22.17***
	[2.943]	[3.788]	[3.837]	[1.984]	[2.115]	[2.116]
Average income	0.297***	0.636***	0.645***	-0.126***	-0.0633	-0.0601
	[0.0848]	[0.139]	[0.141]	[0.0394]	[0.0494]	[0.0496]
City FE	YES	YES	YES	NO	NO	NO
Quarter FE	YES	YES	YES	YES	YES	YES
Time trend	YES	YES	YES	YES	YES	YES
Observations	820	820	820	820	820	820
Adjusted R^2	0.742	0.589	0.579	0.696	0.670	0.668
Test full shifting	1.83	17.12	17.03	1.02	4.93	5.33
	0.1759	0.0000	0.0000	0.3137	0.0265	0.0209
Sargan Test	3.157	6.935	8.036	2.395	7.718	9.590
	0.0756	0.0312	0.0453	0.1217	0.0211	0.0224
Weak identif.test	160.77	13.24	7.93	380.62	18.82	11.34

Intercept included in all regressions.
Standard errors based on robust covariance matrices in brackets in brackets.
Columns 1 and 4: HHI instrumented using all exogenous regressors plus lagged values of HHI (HHI_{t-1} and HHI_{t-2}).
Columns 2 and 5: Excise instrumented using all exogenous regressors plus the share of seats of the political party in power in each province and a dummy indicating whether the political party in power is the same at both the federal and state levels.
Columns 3 and 6: Excise and HHI instrumented using all exogenous regressors plus the share of seats of the political party in power in each province, a dummy indicating whether the political party in power is the same at both the federal and state levels, HHI_{t-1} and HHI_{t-2}.
Test for fully shifting under perfect competition (excise = 1): χ^2 above and P values below.
Sargan test of overidentifying restrictions: χ^2 above and P values below.
Weak identification test: 1st-stage F except for columns 3 and 6 where we provide the Cragg-Donald Wald F statistic to be compared with the values given by Stock and Yogo (2002).
* $p<0.10$, ** $p<0.05$, *** $p<0.01$