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tax shifting based on observables**

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# The Power to Pass on Taxes - A Test for Tax Shifting based on Observables\*

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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 when considering tax incidence 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, this depends on unobservable parameters of the demand and cost functions. In this paper we devise a novel empirical test, based on observables, to assess whether taxes are under- or overshifted in an oligopolistic market. The test depends on the interaction between market structure and taxes. We apply our test to the Canadian retail gasoline market using a panel data set of 10 cities, finding that gasoline taxes are undershifted.

**Keywords:** Tax Incidence; Pass-through; Market Structure

**JEL classification:** H22; D43; L13

<All tables 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, i.e. they can be *under-* or *overshifted*.

The condition whether taxes are under- or overshifted depends theoretically on parameters of the demand and cost functions. These are often unobservable to the econometrician or inherently difficult to estimate. In this paper we devise a novel empirical test, based on observables, to identify whether a market presents under-, full or overshifting. We show that whether a particular market displays under- or overshifting can be assessed by the sign of an interaction term between taxes and market structure. We then take our test to data. 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 identifying 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, an increase in market concentration - measured through changes to city specific Herfindahl-Hirschman Indices -

raises prices. Our main coefficient of interest, the interaction term between the Herfindahl-Hirschman Index and excise taxes, is negative and, in most specifications, statistically significant. This implies that taxes in the Canadian retail gasoline market are undershifted. We apply a series of robustness checks including alternative specifications for inference and controlling for the potential endogeneity of taxes and market structure. Our results are largely confirmed.

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 CO<sub>2</sub> emissions back.<sup>1</sup> 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.<sup>2</sup>

Further, we would argue that our results point to a, potentially important, omission in earlier analysis of tax incidence in oligopolistic markets. First, most empirical applications either implicitly or explicitly assume perfectly competitive markets or they cursory control for market structure (via a main effect). We show that not accounting for the interdependence of market structure and taxes by not including an interaction term might result in a misspecified model. In our study, not accounting for the differential effect of taxes depending on market structure results in an underestimation of tax incidence (evaluated at the mean) of around 20%. Similarly, in our sample, tax incidence varies from not significantly different from full absorption by producers in the most concentrated markets, to not significantly different from full shifting in the least concentrated ones. Thus, policy conclusions might need to be adapted depending on the degree of (local) market concentration. Finally, a point estimate of tax incidence without appropriately controlling for market structure might produce a coefficient larger than but not statistically different from one. Such an estimate might lead authors to conclude that taxes are fully shifted, when indeed they are overshifted. We would argue that

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<sup>1</sup>See, for example, Palazzi (2011).

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

our test is a more precise instrument to assess tax incidence leading to potentially improved policy conclusions.

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 and derives the empirical test formally. 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 and illustrate the potential of inadequate policy conclusions. Finally, in section 7, we provide some concluding remarks.

## 2 Literature Review

Krzyzaniak and Musgrave (1963) were among the first to suggest that in oligopolistic and monopolistic markets taxes could be overshifted to final prices. Subsequently, a number of theoretical papers have taken up this point.<sup>3</sup> Katz and Rosen (1985) show that the Krzyzaniak-Musgrave result can be rationalized in a standard neoclassical model. They find that assuming competitive or monopolistic markets when indeed a sector is imperfectly competitive can lead to bias in the estimation of tax incidence. Further, they illustrate that tax incidence depends on market structure. Similarly, 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, given a certain level of market concentration, a tax hike may indeed increase the producer's equilibrium net price. 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".<sup>4</sup> 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

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<sup>3</sup>A more detailed review can be found in Fullerton and Metcalf (2002).

<sup>4</sup>Seade (1985), p.28.

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. Also, overshifting is more likely under free entry. Delipalla and Keen (1992) compare the incidence of excise and *ad valorem* taxes in two oligopoly models, with and without free entry. They find that *ad valorem* taxes strictly dominate excise ones in welfare terms. Further, they find that under both the Generalized Cournot and Free Entry Oligopoly models, 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 firms *à la* Bertrand. They analyze the effects of tax incidence on prices and profits in both the short and long runs finding, again, that the conditions for overshifting depend on the curvature of the demand function. Interestingly, their results are very close to those of a Cournot model with homogeneous products.

Summarizing these contributions, the pass-through of taxes or the presence of under- or overshifting generally depends on the functional form of the demand and the cost functions, i.e parameters that are not directly observable to the econometrician. Based on the theoretical literature above we derive, in the following section, a test to detect whether a particular market displays under- or overshifting which depends on observables, namely on the interaction between taxes and market structure.

Tax incidence and market structure have also been studied empirically. Early contributions are Harris (1987) and Karp and Perloff (1989). 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 increase 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 find that assuming a competitive Japanese television industry underestimates

consumer tax incidence by around 19%.

Besley and Rosen (1999) examine the incidence of sales taxes for 12 commodities in 155 U.S. cities over the period 1982-1990 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 cannot reject full shifting but for others they find overshifting. Additionally, they find that the response of prices to a variation in taxes is fast, i.e. there is a lag of about one quarter from the tax variation to the observed impact on prices.

Delipalla and O'Donnell (2001) analyze the 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 to consumers. Indeed, after splitting the sample of countries in two different groups, they find that taxes are overshifted in 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."<sup>5</sup> 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 behavior and find that cigarette excise taxes may be differentially shifted across groups of consumers. 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. Bergman and Hansen (2010) study the tax shifting of excise taxes on alcoholic and non-alcoholic beverages using micro-level data from Denmark. They focus on six episodes of tax changes including both increases and tax cuts. They find evidence suggesting large differences across brands and different types of beverages but, in general, tax increases are overshifted whereas tax cuts are undershifted. They interpret their results as evidence of Danish retailers having substantial

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<sup>5</sup>Delipalla and O'Donnell (2001), p.21.

local market power.

Three papers among the large literature on gasoline tax incidence are of particular interest for our study. Chouinard and Perloff (2007) study price differences across U.S. states. They find that most of the increase in national gasoline prices over the 1990s is explained by a rise in the price of crude oil. Thus, the authors find that other factors such as taxes and market power did not have a significant impact on the price increase observed during that decade. On the other hand, differences among states in retail gasoline prices are largely explained by variations in taxes and market power. DiGiacomo et al. (2012) analyze the incidence of excise taxes in fuel markets and exploit their 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, finding evidence supporting that “flexible” taxation schemes focusing 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. They find full pass-through of both state and federal taxes for both gasoline and diesel in most of their specifications. In addition, the authors find that the observed pass-through is immediately translated into final prices. Contrary to our study, none of these papers consider the interplay between the effect of taxes and market structure explicitly and most of them do not even control for market structure. Indeed, some papers implicitly (or explicitly) make the assumption of perfectly competitive markets.

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 base our test on a simple Cournot oligopoly model based on Besley (1989) where  $N$  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 exogenously given, but perform comparative statics. Finally, we concentrate on the effect of excise taxes, the main tax instrument in our dataset.<sup>6</sup>

Let the inverse demand function be:

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

where  $p$  is the consumer 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.<sup>7</sup>

The profit function of firm  $i$  is given by:

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

where  $\tau$  is an excise tax on  $q$  and  $c(q_i)$  is the cost function.

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

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

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<sup>6</sup>The model can be extended in many directions such as allowing for entry into the industry (Besley (1989)) or adding an *ad valorem* tax (Delipalla and Keen (1992)).

<sup>7</sup> $Q_{-i}$  denotes the output of the industry produced by all the other firms but firm  $i$ .

**Assumption 2.**  $c(q) : \mathfrak{R}^+ \rightarrow \mathfrak{R}^+$  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 profit maximizing firm:

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

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

We can re-express (4) as:

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

where, as defined by Seade (1980),  $\eta = \frac{Qp''}{p'}$  is the elasticity of the slope of the inverse demand and  $k = 1 - \frac{c''}{p'}$  represents the relative slopes of the demand and marginal cost curves.<sup>8</sup>

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

$$\eta + N + Nk > 0. \quad (6)$$

Now, given symmetric firms:

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

plugging (7) into the FOCs:

$$p'(Nq)q + p(Nq) - c'(q) = \tau. \quad (8)$$

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<sup>8</sup>We treat  $\eta$  and  $k$  as parameters, abstracting from second-order effects of  $N$  and  $\tau$  around equilibrium.

Differentiating (8) with respect to  $\tau$  and rearranging, we obtain:

$$\frac{\partial q}{\partial \tau} = \frac{1}{p'(\eta + N + k)}, \quad (9)$$

and

$$\frac{\partial Q}{\partial \tau} = \frac{N}{p'(\eta + N + k)}. \quad (10)$$

Therefore,

$$\frac{\partial p}{\partial \tau} = p' \left( \frac{\partial Q}{\partial \tau} \right) = \frac{N}{N + (\eta + k)}. \quad (11)$$

**Proposition 1.** *Under perfect competition ( $N \rightarrow \infty$ ) taxes are fully shifted onto consumers ( $\frac{\partial p}{\partial \tau} \rightarrow 1$ ). As the market gets more concentrated, the degree of shifting ( $\frac{\partial p}{\partial \tau}$ ) moves away from one (1). Indeed, under monopoly ( $N \rightarrow 1$ ), the degree of shifting is the largest in absolute value.*

*Proof.* Inspection of (11) immediately proves Proposition 1. □

Equation (11) confirms that, under perfect competition, taxes are fully passed to consumers. Moreover, it shows that  $|\frac{\partial p}{\partial \tau}| - 1$  is the largest for  $N = 1$ , i.e. under monopoly the degree of shifting is always the furthest away from one. We can now look at the conditions for under- respectively overshifting.

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

*Proof.* Inspection of (11) allows us to classify the cases of relevance for our analysis:

$$\frac{\partial p}{\partial \tau} = \frac{N}{N + (\eta + k)} \begin{cases} < 1 & \Rightarrow \text{undershifting} & (12a) \\ = 1 & \Rightarrow \text{full shifting} & (12b) \\ > 1 & \Rightarrow \text{overshifting} & (12c) \end{cases}$$

□

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

*Proof.*

$$\frac{\partial \left( \frac{\partial p}{\partial \tau} \right)}{\partial N} = \frac{(\eta + k)}{(\eta + N + k)^2} \quad (13)$$

As  $(\eta + N + k)^2 > 0$  for all  $\eta$ ,  $N$  and  $k$ , the sign of (13) equals the sign of  $(\eta + k)$ .  $\square$

**Corollary.** *Let  $HHI$  be an inverse function of  $N$ . Then,*

$$\frac{\partial \left( \frac{\partial p}{\partial \tau} \right)}{\partial HHI} \simeq \frac{-(\eta + k)}{(N + (\eta + k))^2}.$$

Proposition 3 and its Corollary summarize 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. Both  $\eta$  and  $k$  depend on the underlying parameters of the demand and cost functions and are therefore unobservable or inherently difficult to estimate. However, Proposition 3 shows that the sign of the cross-derivative of prices with respect to taxes ( $\tau$ ) and the number of firms ( $N$ ) is equal to the sign of  $(\eta + k)$ . The empirical counterpart of this cross-derivative is an interaction term of taxes and market structure in a regression equation. Now, as is usual in an empirical context, market structure is often characterized by a concentration measure, such as the Herfindahl-Hirschmann Index (HHI), which is an inverse function of  $N$ . The Corollary shows that the interaction term of taxes with a concentration measure indicates the *opposite* sign of  $(\eta + k)$ .

Our framework allows to test whether taxes are under- or overshifted in a particular market based on observables. Indeed, in a price-regression framework the sign of the coefficient of the interaction term between taxes and market concentration will inform about tax incidence. If the interaction term has a negative coefficient taxes are undershifted, while if its sign is positive taxes are overshifted. An empirical application of our test is straightforward and requires variation in taxes and market structure in the relevant product market.

## 4 The gasoline industry and taxes in Canada<sup>9</sup>

The retail 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.

Retailers considerably decreased during the 1990s 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.<sup>10</sup> These stylized facts are consistent with the following observation noted by the Board'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."<sup>11</sup>

Finally, as mentioned in the same report, besides the barriers to entry that are usually present in the gasoline industry (such as economies of scale and capital requirements), costs to both entering and exiting the retail business considerably increased during the 1990s. New environmental regulations were implemented increasing the cost of opening a retail outlet and entry was discouraged by decreasing retail margins throughout most of the decade.

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<sup>9</sup>This section is based on the comprehensive review conducted by The Conference Board of Canada. The report gives an overview of the industry during the 1980s and 1990s covering the sample period of our analysis.

<sup>10</sup>See also Sen and Townley (2010).

<sup>11</sup>The Conference Board of Canada (2001), piii.

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, on average, around 42% of the price. Gasoline taxes in Canada can be divided as follows: federal and provincial excise taxes; the Goods and Services Tax (GST); as well as the Provincial Sales Taxes (PST), where applicable. The federal excise tax is imposed across Canada and added on to the other price components of gasoline (crude oil, refining margin and retail margin). In addition, a similar (excise) tax on gasoline is also levied by provincial governments. Moreover, the GST is levied on all components of the price of gasoline. 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 (PST). The PST is an *ad valorem* tax and is calculated on the total cost of gasoline. Finally, some Canadian cities apply their own transit (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 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.

To illustrate the variation in our raw data we have divided Canada in four geographical regions: Eastern Canada, Quebec, Ontario and Western Canada. Figure 1 illustrates the evolution of (final) retail prices for the four regions. All of them show decreasing retail prices at the beginning of the period and an increasing trend afterwards. Though the inflection year varies for the different regions, i.e. while final prices in Western Canada started increasing in 1992; in Quebec they started augmenting only in 1995. Moreover, taking the two extremes of the time line covered by our sample (1991 and 1997), we observe that final prices in both Eastern Canada and Ontario remained almost constant while they increased in Western Canada and diminished in Quebec.

In the left panel of Figure 2 we display the evolution of market concentration (measured

by the Hirschmann-Herfindahl Index, HHI) across the period covered in our dataset.<sup>12</sup> Interestingly, we observe different patterns. The highest concentration is found in Eastern Canada with HHI values above 2'000, although with a slight decrease of concentration over time. All other regions have witnessed rationalisation over the period, implying increasing concentration levels. The steepest wave of concentration belongs to Quebec, holding the lowest levels of market concentration (HHI below 1400) at the beginning of the 1990s, but ending up with an HHI close to 2'000 by 1997. Both Ontario and Western Canada experienced market concentration but at lesser rates than Quebec. On the right panel of Figure 2 we plot the evolution of excise taxes (in cents per litre) for the four regions. Taxes have generally increased over the period but at differing rates across regions. While Eastern Canada taxed gasoline at the highest levels in 1991, they were overtaken by Quebec and Ontario subsequently. Also noteworthy is that the spread of excise taxes has slightly increased over time by roughly one cent per litre. Figure 3 illustrates the relation between final prices and market concentration (left panel) and excise taxes (right panel). Both variables display a positive effect on prices, as expected. The raw correlation coefficients are 0.27 for market concentration and 0.43 for taxes.

Finally, Figure 4 illustrates graphically our test. We plot again the relation between taxes and prices dividing the sample into low and high concentrated markets.<sup>13</sup> Further, we fit a simple regression line through the scatter plot for each type of market. In the graph, our test corresponds to comparing the slope coefficients of the fitted lines between low and high concentrated markets (indicated in the figure). We observe that the pass-through of taxes to final prices is slightly lower for high concentrated markets (0.76) than for low concentrated ones (0.81). Thus, the effect of taxes on prices decreases with market

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<sup>12</sup>The Hirschmann-Herfindahl Indexes are computed as the sum squared of firm specific market shares and are a common measure of market concentration used by antitrust agencies such as the Federal Trade Commission and the Department of Justice, especially with respect to merger analysis. HHI are computed as the sum of squared individual firm specific market shares. Hence, a monopoly would result in an HHI of 10,000. As detailed in the U.S. Department of Justice's website (<http://www.justice.gov/atr/public/guidelines/hhi.html>), antitrust agencies generally consider markets in which the HHI is between 1,500 and 2,500 points to be moderately concentrated, and markets in which the HHI is in excess of 2,500 points to be highly concentrated.

<sup>13</sup>The sample is divided along the median.

concentration, suggesting that the Canadian retail gasoline market undershifts taxes.

In the following sections, we intend to study the relationship between prices, excise taxes and market concentration using formal econometric models.

## 5 Data and empirical framework

### 5.1 Data

Our dataset consists of monthly observations for the period 1991-1997 of ten of the main cities in Canada.<sup>14</sup> The three main variables are those concerning gasoline prices, taxes and the gasoline retail market structure. We obtained retail prices (the dependent variable) from the Weekly Pump Price Survey (WPPS) of MJ Ervin & Associates.<sup>15</sup> This is a survey of prices for retail gasoline in more than 50 Canadian cities. We use city-level monthly averages and prices are final pump prices including all taxes.<sup>16</sup>

Our main tax variable, *excise*, accounts for federal and provincial excise taxes.<sup>17</sup> Information on market concentration and the number of stations was obtained from The Kent Group, a private consulting company providing this information from all outlets in the markets it surveys. We use the standard Hirschmann-Herfindahl Index (*HHI*) as our market concentration measure.

We include a number of control variables as well. Gasoline wholesale prices and the

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<sup>14</sup>The cities in our database are: Calgary, Halifax, Montreal, Ottawa, Quebec, Saint John, St. Johns, Toronto, Vancouver and Winnipeg.

<sup>15</sup>MJ Ervin & Associates is a division of the Kent Group.

<sup>16</sup>Such taxes include the Goods and Services Tax (GST). Since the GST is also applied to excise taxes, one could argue that using final prices biases our results towards finding overshifting of taxes as any (excise) tax increase would be multiplied by one plus the GST into final prices. We are not particularly worried about this issue, mainly because the GST is a level effect and it is not clear that it should affect our coefficient of interest, namely the interaction of taxes and market structure. We would, indeed, argue that this could be seen as an advantage of our test.

<sup>17</sup>We restrict our analysis to variation of provincial and federal excise taxes as the *ad valorem* sales tax component, whether in terms of GST, PST or HST is a relatively small portion of final per litre retail prices. Indeed, our summary statistics show that excise taxes constitute 62% of final retail prices. Nevertheless, as a robustness check, we also carried out 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) converted into dollar equivalents. Results do not significantly differ and are available upon request.

number of gasoline stations control for supply-side characteristics. The wholesale price is also reported in the WPPS and is highly correlated with crude oil price suggesting that it captures both the effects from the Canadian wholesale and the international crude oil markets. Thus, we decided not to include crude oil price as a regressor. Finally, we also control for demand side aspects by including per capita gasoline consumption, average income at the provincial level and demographic controls (provincial unemployment rates, *population*, percentage of youths between 15 and 24 years old). Table 1 presents summary statistics.

## 5.2 Empirical framework

We estimate the following model:

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

where  $Price_{it}$  denotes retail prices including taxes in city  $i$  at year  $t$ .  $Excise_{it}$  accounts for both federal and provincial excise taxes,  $HHI_{it}$  is the Hirschman-Herfindahl Index and  $\mathbf{X}_{it}$  is the set of control variables described above.<sup>18</sup>

Two comments to this specification regarding our theory are in order. First, equation (11) shows that the effect of taxes on prices, in general, depends on market structure ( $N$  in our model). Second, our test (formulated in Proposition 3), calls for an interaction term between taxes and market structure. Thus, the inclusion of the interaction term in (14) is warranted by theory.

This is relevant, because, to our knowledge, no study of tax incidence in oligopolistic markets has included this specific interaction term ( $\beta_3$ ); not even those that do control for market structure. To say it differently, previous studies on tax incidence measure the effect of taxes on *prices* through the coefficient  $\beta_1$ . Similarly, in these models,  $\beta_2$  measures the effect of market concentration on retail prices. In our case,  $\beta_1$  measures the direct pass-through effect of taxes under perfect competition, i.e. when  $HHI = 0$  and  $\beta_2$  accounts for the direct

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<sup>18</sup>In order to simplify the interpretation of the estimates, we re-scale  $HHI$  between 0 and 1.

effect of market structure on *prices* when no tax is applied. We thus expect  $\beta_1$  to be equal to 1 and  $\beta_2$  to be positive. Finally, our main coefficient of interest is  $\beta_3$ , the coefficient of the interaction term that measures the effect of market concentration on the degree of tax shifting. From Proposition 3 and its Corollary, its empirical sign indicates whether we are in presence of under- or overshifting.

### 5.2.1 Inference

Our basic specifications present 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. Newey-West standard errors allow the error structure to be heteroskedastic and possibly autocorrelated up to some lag. Another strategy in the same direction is to fit the model using Feasible Generalized Least Squares (FGLS). This strategy allows for the presence of AR(1) autocorrelation within panels and cross-sectional correlation and heteroskedasticity across panels.

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), the estimation of the covariance matrix without controlling for clustering can lead to understated standard errors and overstated statistical significance.<sup>19</sup> Under the usual assumptions, the OLS estimator is unbiased in small samples and normally distributed or consistent and approximately normally distributed when we work with a large number of clusters. In other words, the cluster-robust covariance matrix is consistent when the number of clusters tends to infinity. However, we are working with only 10 clusters. Thus the standard procedures for clustering might not be valid. An alternative is the application of the wild bootstrap method applied in Cameron et al. (2008).

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<sup>19</sup>See also Moulton (1986), Wooldridge (2003) and Stock and Watson (2008) among others.

## 5.2.2 Endogeneity

One could argue that two potential issues of endogeneity exist in our analysis. 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. As such, we perform instrumental variables regression.

First, concerning market structure, we can expect  $P_{it}$  and  $HHI_{it}$  to be correlated but it is less likely that  $Price_{it}$  and  $HHI_{it-1}$  or  $HHI_{it-2}$  are. Thus, we use the one- and two-periods lagged value of  $HHI_{it}$  as an instrument for market structure. Similarly, following the strategy implemented by Besley and Case (2000), we instrument taxes with the political party in power at the provincial level. Specifically, we use data on the share of seats of the political party in power in each province and an indicator whether the political party in power is the same at both the federal and provincial levels. The idea behind our strategy is that political parties may influence the set of policies to be implemented (including taxation) without having any direct relationship with retail prices.

# 6 Estimation and results

## 6.1 Preliminary results

We start in Table 2 repeating, with our sample, what has currently been done in the literature, i.e. we estimate (14) without the interaction term. In column (1) we include city-specific fixed effects, while in column (2) we add quarterly dummies to capture seasonality. In column (3) we also include an overall time trend. Finally, in column (4) we omit the city level fixed effects and the overall time trend but allow for city-specific time trends. The overall fit of our model is good with  $R^2$  of around 0.75.

Regarding the control variables, we observe that wholesale prices are an important deter-

minant of retail prices. An increase of the wholesale price of 10 cents raises final prices by roughly 7 cents. Also, the impact of *wholesale* is rather persistent. The lagged value is still positive and highly significant, while the coefficient of the second lag is negative. The socio-economic controls reveal the expected signs. The unemployment rate, an indicator of the economic situation of the city, reduces *prices*, as does a higher share of young. *Population* instead increases *prices*, likely due to increased driving distances and levels of congestion in larger cities. Similarly, *average income* has a positive and significant effect on *prices*.

Of the supply side variables, only the number of stations has a significant effect on *prices*. Slightly contrary to expectations, a higher number of stations raises gas prices. We presume that this result is driven by the fact that larger cities have more stations but also higher demand.

Regarding our main coefficients of interest it is interesting to note that market concentration does not seem to affect retail prices as expected. The coefficient of *HHI* is negative but not statistically significant, suggesting that these specifications might be biased.

More importantly, *excise* exerts a positive and significant (at least at the 10%-level) effect on *prices* across all specifications. The magnitude of the coefficient is rather small. Taking the coefficient of column (3), an increase in *excise* of ten cents raises the final price by roughly 2.5 cents. Evaluated at the mean this implies an elasticity of *prices* with respect to *excise* of 0.07. These results suggest that we are in presence of undershifting and that retail prices react actually quite little to changes in taxes. However, these results do not take into account the possibility that the degree of tax shifting might vary with market concentration, to which we now turn.

## 6.2 Main results

Table 3 shows our main results. The specifications in column (1) to (4) are the same as in Table 2. All the control variables have the identical signs and similar magnitudes. Again, only gasoline consumption is not statistically significant. Hence we will not discuss controls

again and concentrate on our main coefficients.

The direct effect of excise taxes on final prices is, statistically significant and positive in all the specifications. Further, the magnitude of the coefficient varies very little across specifications and is around 1.3. This implies, taking the coefficient values of column (3), that under perfect competition ( $HHI = 0$ ) a 10 cent increase in *excise* is associated with a 13.5 cent raise in *prices*. Thus, the results suggest overshifting of taxes even in competitive markets, which is contrary to theory. However, in none of the specifications we can reject that the main effect of *excise* is actually equal to one, i.e. that taxes are fully shifted when  $HHI = 0$ , as indicated in the last row of Table 3, where we test for  $\beta_1 = 1$ .<sup>20</sup>

Similarly, the direct effect of market structure on *prices* is statistically significant and positive, as expected. Again, the magnitude of the coefficient is robust across specifications and is around 120. Thus, an increase in  $HHI$  of 0.1 implies a raise in *prices* of roughly 12 cents.<sup>21</sup>

Most importantly, the degree of tax shifting depends largely on market structure. Indeed, the coefficient of the interaction term between *excise* and  $HHI$  is negative and highly statistically significant. Again, its magnitude is robust across specifications and fluctuates around  $-6$ . This is quite an important effect. Evaluated at the mean of  $HHI$  ( $= 0.18$ ) our results imply that a 10 cent increase in taxes is associated with a raise of 3.0 cents in final prices, a tremendous difference from the degree of tax shifting under perfect competition (13.5). Incidentally, the average effect of *excise* is slightly higher than the one estimated in the model without the interaction term ( $= 0.25$ ). We return to the interpretation of the results in section 6.4.

Based on our test, the negative sign of the interaction term between taxes and market concentration lets us conclude that excise taxes are undershifted in the Canadian retail gasoline market.

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<sup>20</sup>An interpretation of  $\beta_1 > 1$  could come from the fact that we are estimating out of sample, since we do not observe values of  $HHI$  close to 0 in our sample (see Table 1).

<sup>21</sup>Recall that  $HHI$  in the estimation is re-scaled between 0 and 1. Hence, an increase in  $HHI$  of 0.1 represent a 1000 point increase in the usual scale.

Finally, one could argue that our results are driven by the unit of observation, which is the city-month. This corresponds to the variation we observe for the *HHI*. Thus, we aggregated the weekly prices we avail of. It could be suspected that both market concentration and excise taxes vary little from month to month and that our coefficients are essentially identified on the cross-section dimension. Columns (5) and (6) of Table 2 show that this is not the case. In column (5) we included year fixed effects and a full subset of city-month fixed effects. Our results, do not change qualitatively. However, this is not our preferred specification, as we reject the null hypothesis of  $\beta_1 = 1$ .<sup>22</sup>

Finally, in column (6) we go in the other direction and estimate our model using yearly instead of monthly averages. Again, results do not change qualitatively although we, quite obviously, loose statistical precision.<sup>23</sup>

### 6.3 Robustness checks

We performed a battery of robustness checks of our main results. Table 4 presents alternative specifications of the error structure. Column (1) repeats the results from column (3) in Table 3, which we take as our base specification for the robustness checks. In column (2) we use a random effects model. The next two columns address potential issues of autocorrelation. In column (3) we allow for Newey-West standard errors. Note that in this specification we did not include the lagged values of the wholesale price. Next, in column (4) we estimate a feasible GLS model.

Our results are robust to these alternative specifications and confirm our result that taxes are undershifted. In all of them excise taxes and market concentration exert a positive and significant impact on gasoline prices, while the interaction between these two variables is negative and highly statistically significant. Further, one can note that the magnitude of

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<sup>22</sup>Furthermore, the magnitude of the coefficients of column (5) is larger for all main coefficients. By preferring the specification in column (3) we also take a somewhat conservative approach for discussion of results.

<sup>23</sup>Further, inspection of Figure 1, with the *u*-shaped evolution of prices could call for a break in the time trend. We estimated our model with a break in 1994, allowing for differential trends in these two periods. Results do not change and are available upon request.

the effects is also fairly stable across specifications. The lowest magnitudes are obtained in the random effects model. Finally, in none of the specifications we reject the test of full tax shifting under perfect competition (last row of the table).

In columns (5) and (6) we cluster standard errors at the city-level. As mentioned before, this strategy seems straightforward. In column (5) we use city fixed effects, while in column (6) we use a random effects model.

Note that, when clustering standard errors, we lose statistical significance, except for *excise* in the random effects model.<sup>24</sup> However, as discussed above, given that we work with a very small number of clusters (10), clustering techniques might not be adequate in our situation. Hence, we prefer our main results from Table 3.

In Table 5 we turn to the possible endogeneity issues as described in section 5.2.2. We instrument *excise* with 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 provincial levels. Market structure is instrumented with two lags of *HHI*. The first three columns are based on FE-2SLS whereas columns (4) to (6) replicate the same specifications under a RE-2SLS approach.

In columns (1) and (4), we only instrument *HHI*, while in columns (2) and (5) we only instrument the tax variable, *excise*. Finally, in columns (3) and (6), we tackle both endogeneity issues at the same time instrumenting both *excise* and *HHI*.

Our set of instruments performs quite well in terms of weak instruments. We compute a Weak Identification Test following Stock and Yogo (2002).<sup>25</sup> 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 larger than 10 suggesting that we are not in presence of weak instruments. In columns (3) and (6), where we consider two endogenous variables simultaneously, we provide the Cragg-Donald Wald F-statistic to be

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<sup>24</sup>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.

<sup>25</sup>See also Hausman et al. (2005) and Baum et al. (2003).

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.<sup>26</sup> In column (6), we can reject the null that the set of instruments is weak at a 10% maximal IV relative bias.

Things are slightly more problematic in terms of the exogeneity condition. As it can be observed from Table 5, the Sargan test is passed when we only instrument for market structure. But, the exogeneity test is not passed once we instrument *excise*. We attribute this result to the fact that we have fairly little variation in our instruments for taxes. Not only is the political situation in the different provinces fairly stable over time but, additionally, political variables do not change over the electoral cycle (usually four years in Canada). Thus, the short time period we avail of implies we are not able to identify the coefficients with sufficient precision.

Nevertheless, turning to our main coefficients, we can observe that across all instrumented specifications our main coefficients remain statistically significant and with the expected signs. Hence, instrumenting confirms the result of our test that excise taxes are undershifted in the market under study.<sup>27</sup> Nevertheless, as it is often the case, the FE-2SLS estimates are bigger than the FE-OLS ones. This is particularly the case when we instrument for *excise*. Indeed, when instrumenting for taxes in a fixed effect framework (columns (2) and (3)), our coefficients literally explode while remaining with the expected signs.

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<sup>26</sup>It is worth mentioning 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.

<sup>27</sup>Following the procedure in Baltagi (2009), we perform two Hausman tests based on the contrast between the fixed effects and random effects estimators; one before and other after controlling for endogeneity. The idea is that the result could have changed when applying the 2SLS framework. Nevertheless, in our case, we reject the null that the RE estimator is consistent in both cases, before and after controlling for endogeneity. Hence, the FE specification is preferred to the RE one in both Table 3 and Table 5. We show both the FE and RE estimations as an additional robustness check.

## 6.4 Discussion of results<sup>28</sup>

What do our results imply? In this subsection we briefly discuss four points we consider particularly relevant: i) model specification; ii) policy conclusions; iii) generality of our test; and iv) welfare implications.

We first illustrate our results graphically. Figure 5 plots the level of tax shifting depending on market structure. It can be noted that for a large part of concentration values contained in our sample, tax incidence is statistically significantly less than one.<sup>29</sup> Also, one can infer from the graph that full shifting cannot be rejected for perfectly competitive markets (as we have tested statistically above). Further, the graph nicely illustrates the strong effect market structure has on tax incidence. Indeed, the estimated effect for concentrated markets turns even negative at around  $HHI = 2300$ , although this is never statistically significant.

Let us now turn to our first discussion point. As mentioned above, most empirical applications of tax incidence in oligopolistic markets did not include the interaction term between taxes and market structure, although warranted by theory. These models might be inadequately specified as our results suggest. In our sample, not controlling for the interaction term underestimates tax incidence. In column (3) of Table 2 the tax coefficient was estimated at around 0.25 while, in our preferred specification, tax incidence at the mean of market concentration is 0.30, i.e. an increase of around 20%.

However, not only tax incidence could be incorrectly estimated, also wrong policy conclusions could be taken. This takes us to the next discussion point. Taking the model of Table 2 one would conclude that only about 25% of an increase in taxes is passed on to consumers, and would apply this policy result to all the markets. However, the effective tax incidence varies greatly across different markets (and hence regions). This is again nicely illustrated in Figure 5 above. In our sample, tax incidence varies from 0.65 to (an admittedly implausible)  $-0.44$  between the least and most concentrated markets. More importantly, looking at the inference of our coefficients, our results suggest that in the least concentrated markets (Ot-

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<sup>28</sup>This section is based on our preferred specification i.e. column (3) in tables 3 and 2.

<sup>29</sup>This is the case for  $HHI > 1450$ .

tawa, Montreal, Calgary, Winnipeg and Toronto) taxes might actually be fully passed on to consumers, while they might be fully absorbed by producers in the most concentrated markets (Vancouver, Quebec, Halifax, St. Johns and Saint John). Obviously, policy conclusions might be quite different under these two scenarios.

Further, some caution should be applied when interpreting a model that does not include the interaction term.<sup>30</sup> Suppose that estimating a model without the differential effect of taxes depending on market structure produces a coefficient on *excise* that is not statistically different from one. Then, one would conclude that taxes are fully shifted to consumers, confirming the implicit assumption of a perfectly competitive market.

For example, we think that our findings are relevant to the conclusions drawn by Marion and Muehlegger (2011), in a recent and important contribution to the sparse literature on tax incidence. While the main focus of their paper is not on imperfectly competitive markets, but rather on the effect of capacity constraints on tax incidence, they do estimate price regressions similar to ours for the U.S. gasoline and diesel markets. In the paper, they estimate a total of 40 pass-through rates. In most cases (in 31 out of 40 regressions) they find coefficient estimates greater than 1, but not statistically significantly different from 1. They conclude that "... gasoline and diesel fuel taxes are on average fully... passed on to consumers".<sup>31</sup> However, the authors do not control for market structure. Given this, our findings raise the possibility that taxes may be *overshifted* in the U.S. gasoline and diesel markets. In other words, if one were to estimate a model controlling for market structure and the interaction term between taxes and market structure, one were likely to find a positive interaction term. The model by Marion and Muehlegger (2011) estimates tax incidence essentially at the mean of market structure across States. The confidence interval of this estimation does not allow them to reject full shifting leading to the confirmation of a perfectly competitive market. In Figure 6 we illustrate this alternative scenario. Let's assume taxes are *overshifted*, hence tax incidence increases with market concentration. However, the confidence interval of the pass-

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<sup>30</sup>Although it should be noted that this is not the case in our sample.

<sup>31</sup>Marion and Muehlegger (2011), p. 1202.

through rates estimated at the mean of market concentration might contain one, leading to the potentially incorrect conclusion of full shifting. It should be finally noted that the policy conclusions might be different under overshifting. For instance, consumer surplus reduction is larger (even in case of inelastic demand) and producer profits can increase when taxes are overshifted.

Of course, the above is a possibility which can only be confirmed if access to market structure data in the U.S. markets are available. However, at the very least, we would argue that our test complements Marion and Muehlegger (2011) quite nicely, in terms of being a useful instrument to assess the degree of tax shifting across industries. Indeed, we think that our test can, with reasonable confidence, measure whether taxes are under- or overshifted in a particular industry. This is, in our opinion, an important and relevant policy result.

Third, one can easily imagine a more general application of our test. Essentially, a change in excise taxes is an exogenous (at least for an individual firm) shock to its marginal cost. As such, our theory could in principle go through for any exogenous cost shock to firms. This idea could be of interest to antitrust policy, as our test could identify whether exogenous cost shocks are over- or undershifted to consumers depending on market parameters, rather than say, potential collusion. We intend to study the potential implications for antitrust policy in future research.

Finally, we can turn to the welfare implications of our findings. Besley (1989) discusses the welfare effects of commodity taxation in oligopolistic markets. His Proposition 1 (Besley, 1989, p. 366) states that a marginal increase in excise taxes can be welfare improving if  $E < 0$ .<sup>32</sup> The corresponding expression to Besley's  $E$  in our model is  $(\eta + k)$ .<sup>33</sup> Hence, a necessary condition for marginal tax changes to be welfare improving is that they are passed on less than proportionally to consumers. This is the case in our sample. Thus marginal increases in excise taxes in the gasoline market in Canada might be welfare improving. Again, we argue that this is, particularly in the current policy discussion, an interesting result and

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<sup>32</sup>It should be noted that the Proposition 1 in Besley (1989) seems to present a typographic error that we take into account in the text.

<sup>33</sup>Note that Besley (1989) assumes a constant marginal cost, hence in his model  $k = 1$ .

should be explored further in future research.

## 7 Concluding remarks

The Canadian gasoline retail is a concentrated market. Thus, the perfect competition assumption might be misleading when analyzing the effect of taxes on prices. 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 or overshifting depends on unobservable parameters of the demand and cost functions. In this paper, we devise an empirical test to assess the degree of tax shifting which is based on observables, namely an interaction term between taxes and market structure in a price equation.

We apply our test to the Canadian retail gasoline market. In our sample of monthly observations across 10 cities over the period 1991-1997, we find that taxes are undershifted. The results are remarkably robust to a series of alternative specifications both regarding inference and instrumenting our main variables.

We would argue that our results improve earlier analysis of tax incidence in oligopolistic markets in an important direction. Most of earlier studies in concentrated markets (e.g. gasoline and cigarettes) either explicitly or implicitly assume perfectly competitive markets. We show that such an estimate of tax incidence might be biased when not specifying the model correctly, i.e. when not accounting for the interdependence of market structure and taxes. Further, policy conclusions might be inadequate.

In our data, by not allowing tax incidence to vary with market structure, we find a degree of pass-through of 0.18. Correctly specifying our model implies a tax incidence, evaluated at the mean, of 0.30, i.e. an increase of 67%. We also find that the degree of tax shifting depends heavily on market structure. While for the least concentrated markets we cannot reject that taxes are fully passed on to consumer prices, this result is actually inversed for the most concentrated ones where taxes might be fully absorbed by producers. Quite obviously,

policy conclusions regarding gasoline taxation might be quite different and should depend on the degree of concentration.

Our empirical test is easily applicable to other markets and regions. It requires data showing variation in tax rates and market structure. The Canadian example, where Provinces also set gasoline taxes, is a good illustration of the applicability. The U.S. might be another good example, since States tax gasoline and there is likely also considerable variation in market concentration. Further, our test could have even more general implications and could find its way into further analysis of tax policy and likely even antitrust issues such as collusion. In future work, we aim at linking the vast literature on tax competition with the tax incidence results described in this paper.<sup>34</sup> Another interesting issue to explore is whether and how our results are affected by the presence of different size firms.<sup>35</sup>

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<sup>34</sup>See, for example, Hayashi and Boadway (2001), Devereux et al. (2007) and Brulhart and Jametti (2006) among others.

<sup>35</sup>See Vigneault and Wen (2002).

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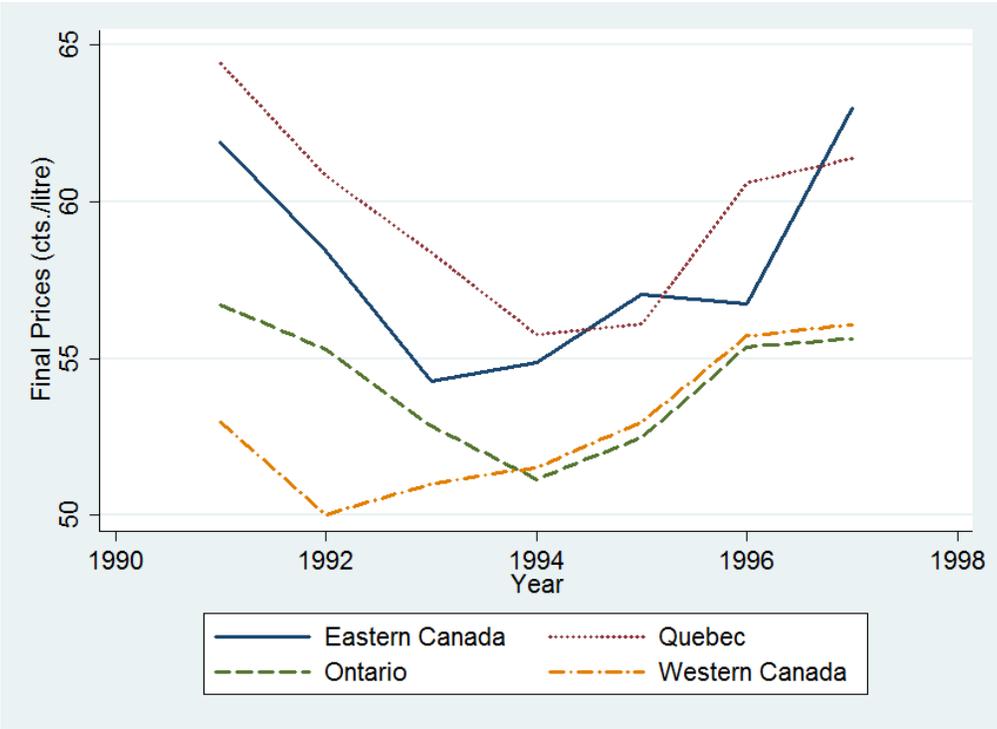


Figure 1: Time Evolution of Prices

Source: The Kent Group

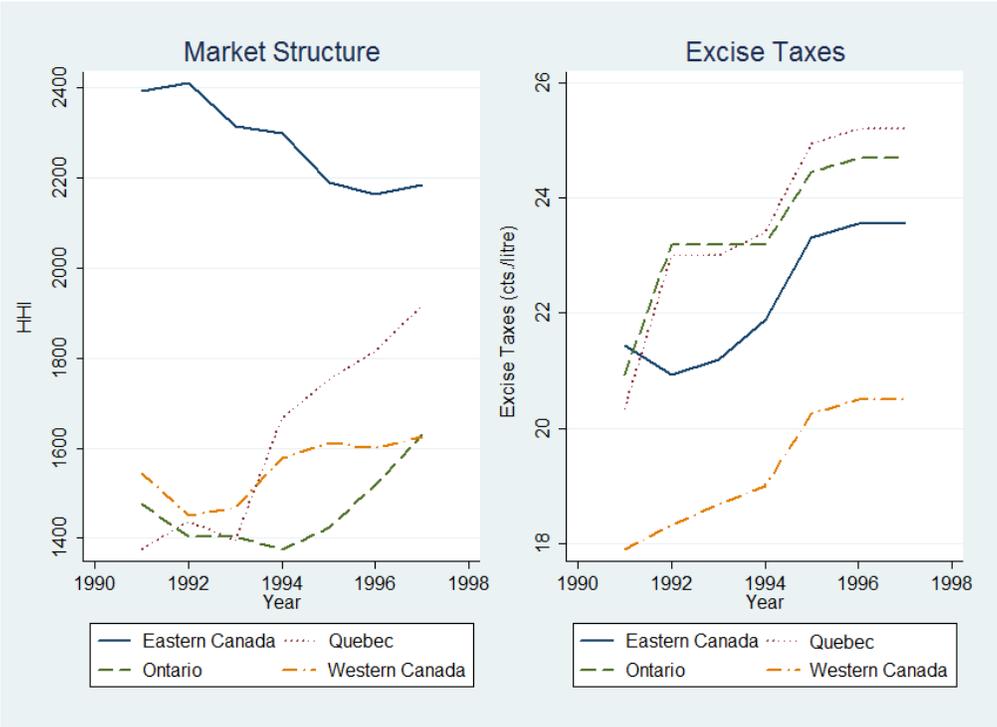


Figure 2: Time Evolution of Market Concentration and Excise taxes

Source: The Kent Group and own calculations



Figure 3: Final Prices, Excise Taxes and Market Structure

Source: The Kent Group and own calculations

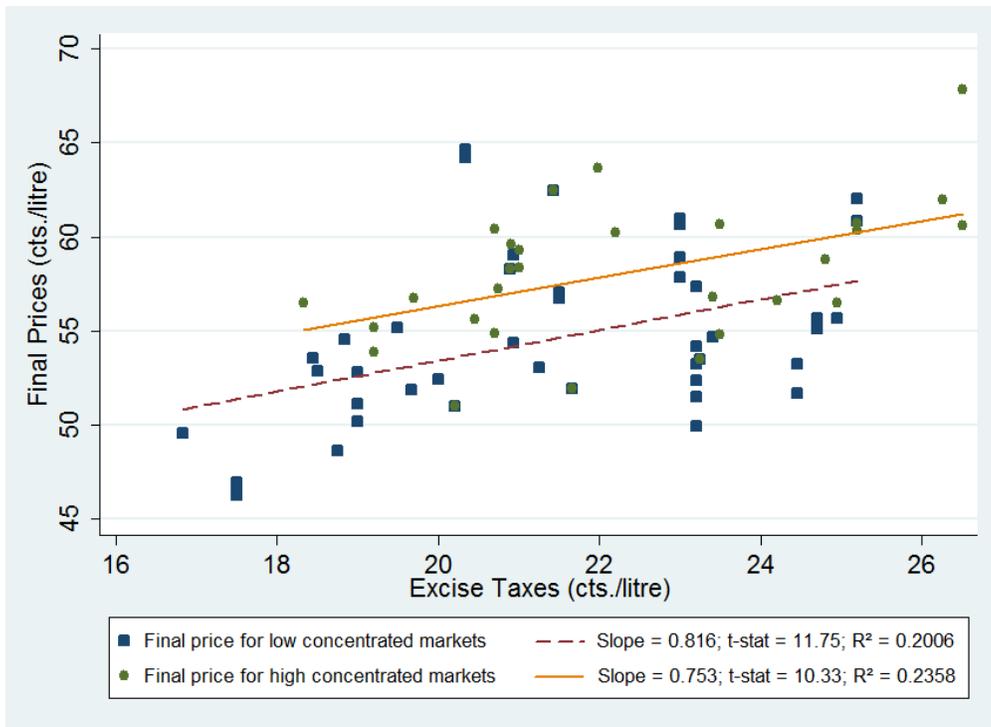


Figure 4: Prices and Taxes for Low and Highly Concentrated Markets

Source: The Kent Group and own calculations

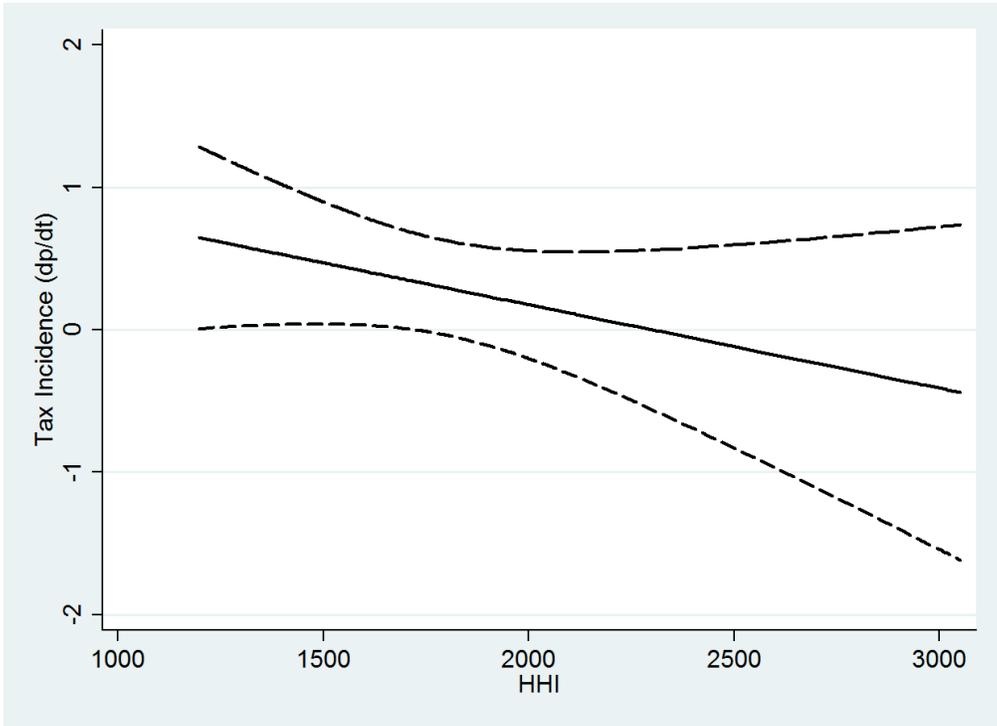


Figure 5: Interdependence of Tax Incidence and Market Structure

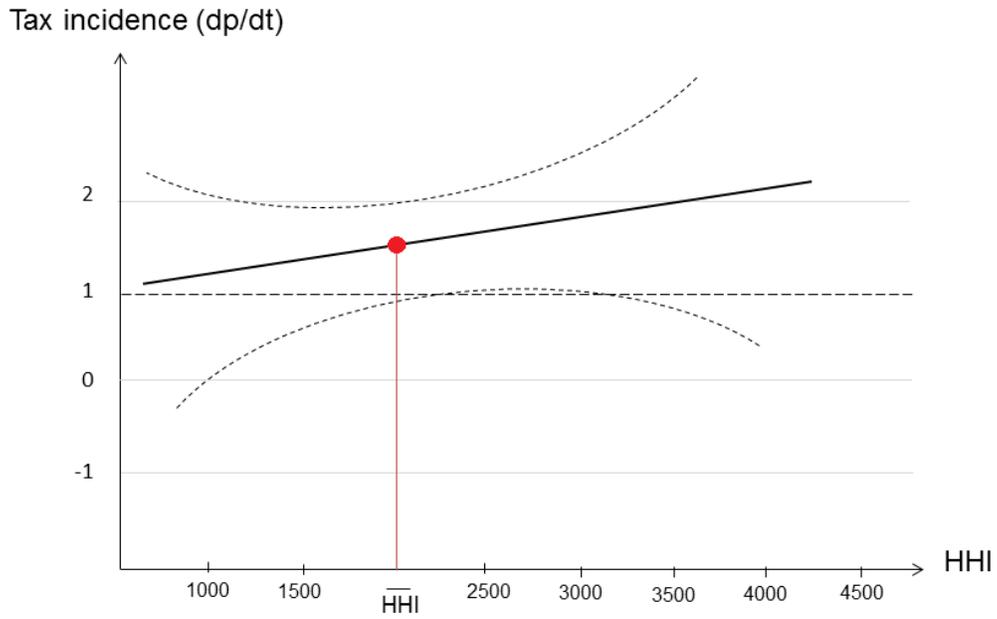


Figure 6: An Alternative Scenario

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 Hirschman 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
Share of seats pol. party in power	%	840	65.71	9.46	52	87.27
Both lib. in power	Dummy	840	0.32	0.47	0	1
Both conserv. in power	Dummy	840	0.60	0.24	0	1

Table 2: No Interaction Term - OLS and Fixed Effects Estimates

	(1)	(2)	(3)	(4)
Excise	0.180* [0.0954]	0.188** [0.0955]	0.256** [0.124]	0.182* [0.0957]
HHI	-5.199 [7.802]	-5.205 [7.745]	-5.798 [7.661]	-5.107 [7.743]
Wholesale	0.669*** [0.0734]	0.640*** [0.0785]	0.644*** [0.0791]	0.639*** [0.0785]
Wholesale_lag1	0.562*** [0.111]	0.540*** [0.113]	0.541*** [0.113]	0.540*** [0.113]
Wholesale_lag2	-0.241*** [0.0847]	-0.225*** [0.0833]	-0.229*** [0.0833]	-0.225*** [0.0833]
Unemployment rate	-0.311*** [0.116]	-0.326*** [0.114]	-0.356*** [0.120]	-0.324*** [0.114]
Population	0.488** [0.192]	0.542*** [0.193]	0.571*** [0.192]	0.555*** [0.195]
Prop. aged 15 to 24	-0.221* [0.131]	-0.218* [0.130]	-0.264* [0.136]	-0.214* [0.130]
Stations	9.856*** [3.170]	10.54*** [3.250]	9.698*** [3.413]	10.40*** [3.192]
Gasoline cons. pp	-0.602 [3.062]	-1.785 [3.009]	-1.493 [3.034]	-1.866 [3.009]
Average income	0.226** [0.0888]	0.235*** [0.0866]	0.236*** [0.0866]	0.236*** [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
$R^2$	0.7433	0.7468	0.7471	0.7467

Intercept included in all regressions. Standard errors based on robust covariance matrices in brackets.  $R^2$  overall where applicable.

\*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$

Table 3: Main Results - OLS and Fixed Effects Estimates

	(1)	(2)	(3)	(4)	(5)	(6)
Excise	1.361*** [0.343]	1.304*** [0.340]	1.348*** [0.357]	1.311*** [0.340]	1.772*** [0.384]	2.752** [1.297]
HHI	127.6*** [39.18]	120.4*** [38.96]	118.5*** [38.69]	121.9*** [38.93]	134.1*** [41.34]	228.0* [122.2]
Excise_HHI	-6.264*** [1.793]	-5.924*** [1.793]	-5.859*** [1.782]	-5.993*** [1.793]	-6.934*** [1.908]	-10.21* [5.806]
Wholesale	0.691*** [0.0732]	0.663*** [0.0784]	0.666*** [0.0790]	0.663*** [0.0783]	0.504*** [0.0943]	1.484 [0.937]
Wholesale_lag1	0.547*** [0.111]	0.527*** [0.113]	0.527*** [0.113]	0.527*** [0.113]	0.552*** [0.126]	-0.355 [0.727]
Wholesale_lag2	-0.204** [0.0875]	-0.191** [0.0858]	-0.194** [0.0856]	-0.190** [0.0858]	-0.320*** [0.0891]	
Unemployment rate	-0.241** [0.120]	-0.259** [0.118]	-0.284** [0.123]	-0.256** [0.118]	-0.432*** [0.153]	0.0987 [0.809]
Population	0.462** [0.198]	0.516*** [0.199]	0.540*** [0.197]	0.529*** [0.201]	0.477** [0.204]	-2.197 [4.996]
Prop. aged 15 to 24	-0.271** [0.134]	-0.265** [0.133]	-0.302** [0.139]	-0.261** [0.133]	-0.323** [0.147]	-1.014 [0.743]
Stations	11.17*** [3.270]	11.76*** [3.324]	11.05*** [3.463]	11.63*** [3.264]	9.068*** [3.252]	7.000 [9.637]
Gasoline cons. pp	-1.028 [3.012]	-2.124 [2.969]	-1.880 [2.992]	-2.200 [2.968]	-1.594 [3.284]	54.81 [179.1]
Average income	0.274*** [0.0872]	0.281*** [0.0856]	0.281*** [0.0855]	0.282*** [0.0856]	0.216** [0.100]	-0.132 [0.371]
City FE	YES	YES	YES	NO	NO	YES
Quarter FE	NO	YES	YES	YES	NO	NO
Year FE	NO	NO	NO	NO	YES	YES
City-Month FE	NO	NO	NO	NO	YES	NO
Time trend	NO	NO	YES	NO	NO	NO
City Specific Time trend	NO	NO	NO	YES	NO	NO
Observations	820	820	820	820	820	60
$R^2$	0.7474	0.7505	0.7507	0.7505	0.7923	0.8976
Test full shifting	1.10 0.2933	0.80 0.3717	0.95 0.3297	0.83 0.3613	4.04 0.0444	1.83 0.1765

Intercept included in all regressions. Standard errors based on robust covariance matrices in brackets.  $R^2$  overall where applicable.

Columns 1 to 5, based on monthly data. Column 6, based on yearly data.

Test for full shifting under perfect competition (excise = 1):  $\chi^2$  statistic above and P values below.

\*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

Table 4: Robustness Checks - Inference

	(1)	(2)	(3)	(4)	(5)	(6)
Excise	1.348*** [0.357]	1.187*** [0.218]	1.209** [0.475]	1.469*** [0.324]	1.348 [0.847]	1.187* [0.676]
HHI	118.5*** [38.69]	93.04*** [24.37]	106.2** [53.81]	131.5*** [37.37]	118.5 [83.40]	93.04 [72.66]
Excise_HHI	-5.859*** [1.782]	-3.871*** [1.229]	-5.224** [2.471]	-6.422*** [1.702]	-5.859 [4.608]	-3.871 [3.934]
Covariates	YES	YES	YES	YES	YES	YES
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
$R^2$	0.7507	0.7025			0.497	
Test full shifting	0.95 0.3297	0.74 0.3904	0.19 0.6594	2.10 0.1474	0.17 0.6906	0.08 0.7822

Intercept included in all regressions. Standard errors in brackets.  $R^2$  overall where applicable.

Column 1 reproduces column 3 in Table 3, our main specification.

Column 2, based on robust covariance matrices.

Column 3, based on Newey-West standard errors. 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.

Column 4, feasible GLS.

Columns 5, FE, clustered standard errors at the city level.

Columns 6, RE, clustered standard errors at the city level.

Test for full shifting under perfect competition (excise = 1):  $\chi^2$  (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

	(1)	(2)	(3)	(4)	(5)	(6)
Excise	1.797*** [0.589]	8.830*** [1.893]	9.088*** [1.960]	1.319*** [0.316]	3.249*** [1.013]	3.345*** [1.016]
HHI	171.7** [66.28]	867.6*** [195.6]	897.7*** [204.2]	108.5*** [36.18]	314.2*** [108.9]	324.5*** [109.2]
Excise_HHI	-8.281*** [3.018]	-41.32*** [9.119]	-42.70*** [9.499]	-4.635** [1.805]	-14.78*** [5.376]	-15.28*** [5.392]
Covariates	YES	YES	YES	YES	YES	YES
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
$R^2$	0.742	0.589	0.579	0.696	0.670	0.668
Test full shifting	1.83 0.1759	17.12 0.0000	17.03 0.0000	1.02 0.3137	4.93 0.0265	5.33 0.0209
Sargan Test	3.157 0.0756	6.935 0.0312	8.036 0.0453	2.395 0.1217	7.718 0.0211	9.590 0.0224
Weak identif.test	160.77	13.24	7.93	380.62	18.82	11.34

Intercept included in all regressions. Standard errors in brackets in brackets.  $R^2$  overall where applicable. Columns 1, 2 and 3, with FE-2SLS. Columns 4, 5 and 6, with RE-2SLS.

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: Both excise and HHI instrumented using the instruments mentioned above.

Test for fully shifting under perfect competition (excise = 1):  $\chi^2$  above and P values below.

Sargan test of overidentifying restrictions:  $\chi^2$  above and P values below.

Weak identification test: 1st-stage F-statistic 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

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