A ‘Long March’ Perspective on Tobacco Use: 

The Rational Addiction Model Revisited*

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February 2003

Abstract

This paper revisits the rational addiction model of cigarette demand. Methodologically, we propose that virtually all tobacco demand studies in the time-series mode suffer from a misspecification of the dependent variable, and second that the presence of lead and lag consumption terms as regressors will generally require more than an instrumental variables estimator to insure consistency. Using Canadian data we find that government regulation has had a strong deterrent effect over the period 1972-2000, and that the price elasticity of demand is now lower than even the recently-obtained low estimates propose. These findings have strong public-policy content.

JEL Classification: I12, C13, C22.

*Acknowledgment: The authors would like to thank SSHRC for financial support. An earlier version of the paper benefited from comments by the participants at the CEA 2002 meeting, Kevin Milligan and Anindya Sen.

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1 Introduction and Background

The objective of this paper is to examine the effectiveness of different measures which governments can use to deter smoking. In particular, is it primarily the bundle of regulations that have been imposed on consumers in recent decades that have been responsible for reducing consumption, or have tax-induced price increases been a more effective method of control? While there have been an enormous number of tobacco/cigarette demand studies since Becker, Grossman and Murphy’s pioneering work (1988, 1994) on rational addiction (RA), both in a time-series and panel-data framework, we focus upon some methodological issues that have remained unaddressed. Specifically, the definition of the dependent variable ‘cigarettes’, in virtually every time series study, remains problematic and has led to overestimates of price elasticity. Second, in estimating the RA model, the lead and lag terms of the dependent variable which appear on the right hand side of the estimating equation demand specific treatment: the standard method is to use instrumental variables. However, some recent studies (e.g. Auld and Grootendorst, 2002) point to pitfalls in estimating rational addiction models of tobacco demand using aggregate data. The problem becomes especially acute in the presence of possible nonstationarity and cointegration among the variables of interest. Kitamura and Phillips (1997) proposed extensions of the instrumental variables method to nonstationary regressions but the structure of the rational addiction model appears to be in conflict with some of their conditions for consistent estimation.

We confront these issues by adopting an econometric framework that accounts explicitly for nonstationarity of the variables and possible parameter instability in the model. We suggest several reparameterizations of the rational addiction model to obtain robust estimates of the long-run and short-run elasticities of cigarette consumption, and study the public policy implications of its sensitivity to price changes using Canadian data. Canada, like most economies, has experienced more extensive tobacco control mechanisms in the past two decades, and these have been directed in many instances towards youth. Non-price measures, in the form of work-place and public-place tobacco use are now widespread. Grim warnings on the health consequences of tobacco use have become mandatory on cigarette packaging. In the area of taxation there is a recent degree of convergence in provincial and federal policy across the provinces. Tax policy in the early- and mid-
nineties was volatile and led to massive variations in tobacco prices, both in Canada as a whole during the period and between the provinces at any given time. In 2001 and 2002, the federal government moved with the provincial governments to increase the price of tobacco in those provinces where it had been relatively low, thereby making the price somewhat more uniform across the economy. But the very considerable price variation experienced in the last two decades provides a fertile ground for investigation.

Several studies investigated the price sensitivity of tobacco use in the seventies and eighties, and a small number have used the most recent period. Reinhardt and Giles (2001) focus on data only up to 1990. We will show that the most recent period is critical for current policy, insofar as our rolling estimates of the various elasticities indicate that consumers are now much less sensitive to price variations than heretofore. Galbraith and Kaiserman (1997) addressed, *inter alia*, the data problems of the early nineties, and they included exports in their consumption base, as a proxy for non-legal domestic sales of tobacco. Their principal findings were that, on a national level, the income elasticity of demand is positive and less than unity, and the price-elasticity is in the neighborhood of -0.4. Auld and Grootendorst (2002) focus primarily upon testing false acceptances of the RA model. Balgati and Griffin (2001) reexamine the choice of panel-data estimators and obtain results that are more supportive to the RA hypothesis. Gruber, Sen and Stabile’s (GSS, 2002) primary contribution comes in the form of using survey data, in addition to time series data, in order to address the data deficiencies resulting from smuggling in the nineties. Their results yield price elasticity estimates in the $\{-0.4, -0.5\}$ range.

The estimation of demand behavior within a time-series framework warrants continued examination for a number of reasons. The first is methodological, and is ignored almost universally. Cigarettes come in essentially two forms: manufactured cigarettes which account for the bulk of sales, and fine-cut tobacco - frequently referred to as ‘hand-rolled’ or ‘sticks’, which accounts for a smaller share of the market. Each product delivers the same utility-enhancing toxins. Canadian consumers have shown a high degree of substi-

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1Following the tax increase of 1991, tobacco smuggling became a major problem. Domestically produced tobacco was exported to the US and subsequently reimported illegally, mainly in southwestern Quebec and southeastern Ontario across the St Laurence river. By late 1993 it is estimated that the illegal sales of such re-imported product was accounting for one half of all sales in the central provinces of Canada.
tution in their preferences for these two products. For example, in 1999, consumers in higher-tax provinces (Manitoba westwards and Newfoundland) consumed fine-cut tobacco at almost twice the rate of consumers in the remaining, lower-tax, provinces, as illustrated in Table 1. When relative prices displayed less variation across the provinces - as in the eighties, such differential consumption patterns were less pronounced. This is shown in Table 2, where the ratio of fine-cut to total sales in the four Western Provinces is much lower than in 1999. Despite the considerable importance of this market, and its variability, it is remarkable that so few econometric studies include both types of cigarette product in their demand studies. Fine-cut tobacco has accounted for between 10% and 15% of the market in the modern era. But if consumers switch to and from this product as a result of changes in its relative price, and as a result of price changes in tobacco in the aggregate, measurement error will characterize models which are based upon a variable that excludes some of the quantity. Chaloupka and Warner, in their very extensive survey of smoking (2000), point to the existence of just a few studies that examine products other than manufactured cigarettes, and these deal primarily with smokeless tobacco. But there are virtually no recent studies that address the demand for fine-cut cigarettes. This is a non-trivial omission given its share of the market.

Furthermore, it has been widely documented that many aggregate time series, including consumption, prices and income, are characterized by stochastic trend behavior. Some interesting methodological issues arise in the RA model of Becker, Grossman and Murphy (1988, 1994) from the simultaneous presence of nonstationary regressors, lead and lag levels of the dependent variable, endogeneity and autocorrelated errors. We address this problem by reparameterizing the model and separating the regressors into stationary and nonstationary components. The new parameters on the nonstationary components have the interpretation of long-run elasticities and their super-consistency is unaffected by the presence of endogenous stationary components. We also examine the possibility of parameter instability in the model over time and construct recursive and rolling sample estimates of price and income elasticities. Our empirical findings of reduced price sensitivity in the last part of the sample reinforce the conjecture that price and tax policies may have much more limited impact on aggregate demand than previously thought.

In the second instance there has been a renewed interest in the discipline on both the capacity and
appropriateness of government regulation of the consumption of tobacco and addictive substance more generally. Practically all economies have experienced in recent times a widened menu of regulations governing consumption in schools, workplaces, public places, restaurants etc. While the RA model might claim that such regulation is not welfare improving, other strains of thought offer a different perspective. Three distinct alternatives are discernible in the recent literature. The first is attributable to Laibson and his co-authors (e.g. Laibson, 1997). In this world individuals have hyperbolic discount functions which evaluate the relative importance of consumption in adjoining periods in a way that depends upon how far in the future these periods are. In essence this is dynamically inconsistent planning. Hence government action can be welfare improving - by imposing constraints on excessive consumption today: Gruber and Koszegi (2001) in a very novel approach, develop the idea that corrective taxes on ‘internalities’ may be appropriate. A second approach is that individuals may suffer from ‘projection bias’ (see O’Donohue and Rabin, 1999). Such bias can come in a number of forms, but is based upon an incorrect specification in the present either of future tastes or future effects of current choices. The third, and most recent, line of investigation is due to Bernheim and Rangel (2002) who frame their approach in psychological terms. Their key premise is that the neocortex may suffer from what they call characterization failure, as a result of stimuli or cues. Such mischaracterization of decisions may lead individuals to make choices they would not otherwise make and may lower life-cycle utility over what a correct characterization of choices would induce. While they discuss cues in terms of adverse stimuli, i.e. stimuli which more incline the individual towards drug consumption, cues can equally be considered as messages or regulations which counteract such tendencies - for example health warnings on packages.

In sum, there are several standpoints which validate a welfare-improving role for regulation, and therefore it is important to ascertain if the restrictive policies which have come into being have been effective. While these insights have not yielded directives on appropriate econometric specification, in contrast to the RA model of Becker, Grossman and Murphy (1988, 1994), it should be borne in mind that the latter takes shape as a result of imposing a quadratic utility function, which in turn gives rise to linear decision rules.

Third, we postulate that the changing composition of the Canadian population during the period of
examination may have played a significant role in the long-term decline in usage. Canadian immigration patterns have changed radically since the sixties, away from European-source countries and towards Asia, Africa and the Caribbean. Such populations have significantly lower smoking rates. Health Canada reports in the *National Population Health Survey* that non-European immigrants have a smoking prevalence in the region of just 10%, in contrast to native-born Canadians whose rate approaches 25%-30%. Furthermore, while ‘assimilation’ may result in the children of such immigrants adopting the same smoking patterns as native-borns, the fact that smoking initiation almost invariably takes place in the teen years, means that adult migrants are unlikely to become smokers. This is born out in the data: recent migrants of non-European source have similar prevalence rates to long-resident immigrants. To our knowledge, no study of tobacco demand has examined the possible extent of this demographic trend.

Fourth, quite remarkable patterns are beginning to emerge from 2001/02 sales data - which are outside of our sample. Tax changes resulted in a real increase in the price of cigarettes of about 35% between the first half of 2000 and the first half of 2002. Shipments of manufactured cigarettes in the domestic market declined by only 7.7% however, and sales of fine cut appear to have been affected to an even lesser degree. This suggests that aggregate demand elasticity in the short to medium term may be even less than what econometricians have previously reported. Our results clearly show a changing time profile of aggregate elasticity and significantly lower price sensitivity in the second half of nineties. Such low values have implications for public policy which are quite different from what behavior in earlier decades would have dictated.

Fifth, we use a sample period that we believe is informative. Six additional years of data are available in the nineties beyond that used in Galbraith-Kaiserman (GK), whose data end in 1994 - the point when the national price fell dramatically. With a product that has a strong habit component, the effect of such a change could take time before becoming manifest in smoking patterns. Indeed, as is evident in our figure 1, there was a significant immediate increase in total cigarette sales in 1994/95, followed by a steady decrease thereafter. This fall in consumption rates surprised many analysts, who believed that the tax reduction of 1994 would spell disaster for longer-term smoking rates. In contrast, GSS’s data run to 1999 and also have
the advantage of being province-specific; this enables them to exclude provinces where smuggling was more problematic. However, like GK, the GSS data start in 1980/81, a period that marked a peak in per-capita cigarette use in Canada in the modern era. Consumption rose steadily in the seventies but declined even more rapidly in the eighties (see Figure 1). Since data are available for earlier years, we exploit these. Critical end points in time series may lead to uncertain inference where either time trends or particular error structures characterize the data.

Finally, we note that the nature of the manufactured cigarette has changed significantly in the post-war period, in that it now contains about half of the tobacco weight that it had in 1950, and a progressively lower potential for the delivery of tar and nicotine. While the greater part of the weight reduction took place in the fifties and sixties, the period 1970-2000 still witnessed a moderate decline. Kaiserman and Leahy’s calculations (1992) indicate that tobacco per cigarette declined from 1.6 grams in 1950, to 1.2 grams in 1960, to 0.9 grams in 1970 - the starting point for our data. Since that time, the rate of decline in tobacco per cigarette has been smaller: the typical cigarette now contains about 0.75 grams of tobacco - a reduction of 17% over the period of analysis. This content decline reduces the delivery potential of the utility-enhancing toxins that are craved by the product’s consumers. In conjunction, some measures of tar and nicotine per cigarette have fallen dramatically, as a result of the spread of ‘light’ and ‘mild’ brands.

However, we are reluctant to make any correction to the data to account for these developments: the epidemiological evidence indicates that consumers extract a greater percentage of the potential tar and nicotine content from the ‘low-yield’ variant than from the regular strength product (Jarvis et al., 2001). The ‘mild’ and ‘low-tar/nicotine’ brands achieve their goal primarily by providing a more porous sleeve through which the toxins can escape before entering the body. Yet smokers tend to block the escape route afforded by the increased porosity with their mouth or fingers, and they also tend to draw with greater frequency and vigor than when smoking a ‘regular’ product. The actual toxin-delivery potential of the tobacco in brands that are labelled ‘light’ or ‘mild’ differs rather little from the ‘regular’ strength product. The tar-intake measures listed on packages are those resulting from machine-smoked cigarette experiments as opposed to cigarettes smoked by humans. When the saliva of smokers is examined, the difference in
nicotine content between smokers of regular and light/mild products is small (see Jarvis et al., 2001; and Kozlowski et al., 1998).

A slightly different perspective is taken by Evans and Farrelly (1998), who propose that consumers may switch towards brands that are heavier in tar and nicotine in response to tax increases that are equal per cigarette type. In view of the difficulty that would be associated with defining a cigarette as a weight of tobacco or in terms of toxin delivery, we use the tax authorities’ de facto definition - a sleeve containing tobacco, perhaps having varying weight over time and varying toxin-delivery potential.

The paper is developed as follows: in section 2, trends and patterns in the tobacco market for the recent period are presented. Data issues and priors are addressed in section 3. In view of the very extensive recent review of tobacco demand by Chaloupka and Warner (2000), we omit any further review of the demand literature. We also explore if the growing regulatory environment may have had an impact on the value of the demand elasticity over the period. The modeling framework and estimation procedures are discussed in section 4, and results are presented in section 5. The main conclusions are drawn together in the final section and the results are discussed from the standpoint of public policy.

2 Consumption Patterns

The stylized facts for the Canadian tobacco market in recent times are these:

1. The total number of cigarettes smoked per person reached a peak in the period 1980-82. Thereafter, consumption has fallen secularly. This is illustrated in Figure 1, where the total (seasonally adjusted) sales of manufactured cigarettes per person are plotted. Of the decline since 1980, most has come in the form of reduced prevalence, rather than reduced quantity per smoker².

²The tendency for cigarettes smoked per smoker to remain constant in the face of a declining prevalence springs from the pattern of quitters: those who smoke less are more likely to quit, leaving a smaller number of heavier smokers. Prevalence rates have been tracked by Gilmour (2000), who details the findings from a number of smoking surveys between the late eighties and late nineties. While under-reporting is invariably a problem with surveys dealing with the consumption of toxins, the prevalence reduction that he reports is so strong that it cannot be attributed simply to greater under-reporting in more recent samples. Irvine (2001) has compared survey results with sales figures as a means of validating the reliability of surveys. He found that,
2. The real price of tobacco increased dramatically in the 1980s as a result of tax rate increases (see Figure 2). The greater part of the ensuing revenue growth accrued to the provincial governments. Further tax increases in 1989 and 1991 ushered in a period of chaos. Canadian-produced tobacco was exported and re-imported illegally on a massive scale. By 1994 over one third of national tobacco consumption was in the form of illegal re-imports.

3. In 1994 the federal government tackled the smuggling problem by offering to the provinces a choice of tobacco-taxation regimes: the provinces that agreed to lower their levies would have a lower rate of federal tax in their jurisdictions, while the provinces choosing to maintain their taxes at a higher level would see the federal government do likewise. While this policy basically resolved the problem of smuggling from the US, it led to some inter-provincial movement of tobacco from low- to high-tax jurisdictions. The amounts so moved appear to have been modest in view of the incentive provided by the price differential (Irvine, 2001). In the latter half of the nineties some provinces implemented tax increases. Finally, in 2001 and 2002, most of the remaining low-tax provinces increased their taxes by a significant amount, in conjunction with federal rate increases. The real price of cigarettes at the end of 2001 was almost at the level it attained in early 1994.

This price increase represented a significant change in direction on the part of the federal government. Prior to that time Ottawa feared a renewed bout of smuggling would be triggered by a domestic tax increase. However, cigarette prices in the US rose very significantly in the late nineties. Such increases reflected primarily producer-driven changes rather than higher tax rates. The higher producer prices were considered necessary to cover the envisaged costs of legal settlements with states and consumers (or their surviving relatives) by the tobacco companies (Gruber, 2001).

4. Throughout the period of investigation (1972-2000), but more significantly since 1980, the industry has been the subject of increased governmental regulation. Restrictions on the use of tobacco in the workplace have been implemented by provincial governments; health warnings on cigarette packaging have become mandatory; restrictions on advertizing, marketing and promotional activities have become more prevalent.

5. Total sales, sales per capita and prevalence rates in the year 2001 attained their lowest level in the while not all surveys are equally good, the response variability is moderate.
3 Data, Variables and Priors

All data are collected at quarterly frequency and cover the period 1972:Q1-2000:Q4. The principal source for the various series is the Canadian Socio-economic Information and Management Database (CANSIM). All series are seasonally adjusted using the additive X11 seasonal filter of the U.S. Bureau of the Census.

**Quantity.** Our treatment of the cigarette aggregate is determined primarily by the availability of information on the price side. Statistics Canada furnishes a price series for (manufactured) cigarettes, and one for all tobacco products. The latter is appropriate for our definition of cigarettes, broadly defined. No price series is available only for fine-cut. Accordingly we aggregate the two types of cigarette by means of a constant elasticity of substitution aggregator (utility) function:

\[
C = \left[ \alpha^{\frac{1}{\sigma}} c_{fc}^{\frac{1}{\sigma-1}} + (1 - \alpha)^{\frac{1}{\sigma}} c_{m}^{\frac{1}{\sigma-1}} \right]^{\frac{\sigma}{\sigma-1}}.
\]

The subscripts are self explanatory, \(\alpha\) is a distribution parameter and \(\sigma\) is the elasticity of substitution. Demand functions at the lower level are not used here, but they illustrate that the expenditure shares for each component \(p_{ic_{i}}/M\) may be written as

\[
\frac{p_{ic_{i}}}{M} = \frac{\alpha_{i}p_{i}^{1-\sigma}}{\sum_{j} \alpha_{j}p_{j}^{1-\sigma}}.
\]

Empirically, the share of tobacco expenditure on fine-cut tends be about 10\%, and the price of fine-cut hovers around 70\% of the price of manufactured cigarettes for the period under study. Furthermore, to parameterize \(\sigma\) we can use price and share variation of the products in Quebec in the late nineties. These suggest a value in the region of 1.5 for \(\sigma\) and therefore values \(\alpha_{fc} \simeq 0.07, \alpha_{m} \simeq 0.93^{3}\).

The total manufactured cigarette sales are the sum of domestic sales of cigarettes (CANSIM label D2091)

\[\text{In leaving this we note that there exists an exact price index for this aggregate given that we have been willing to define a quantity aggregator function. That is, in order that the sum of expenditures on the two components equal total expenditure } P_{c}, \text{ since } C \text{ has an explicit form so too does the price index } P_{c}. \text{ In theory, then we could retrieve a price index for fine-cut if we have an index for the aggregate and for the other component. However, this might be too ‘heroic’ a step to take.}\]
and cigarette sales to embassies and for export (CANSIM label D2095). The corresponding fine cut tobacco series (CANSIM label D2093) is converted from kilograms into millions of equivalent cigarettes.

It should be noted that the sales data are shipments from the producer to the wholesaler, not expenditures by consumers. The average time for a pack of cigarettes to be sold to the consumer is several weeks implying that the quantity variable may not be synchronous with the explanatory variables at higher frequency data.

**Price.** The cigarette price index (CANSIM label P100267) and all-items consumer price index (CANSIM label P100000) are used to construct a real price index for cigarettes.

**Income.** The series for the real per capita disposable income is obtained by deflating the disposable income series (CANSIM label D15743) by the GDP deflator and total population (CANSIM label D1). The dependent variable is also deflated by the population. Income per capita is the standard variable used in this area, even though a superior measure might be the income growth of the different quartiles of the income distribution, since it is well recognized that prevalence differs by education and income. We also know that the lowest quartile of the earnings distribution experienced a significantly poorer growth in the period from the late seventies to the mid nineties than did the other quartiles of the distribution.

**Time trend.** In addition to the conventional ‘unmodelable behavior’ rationale, there is one major determinant of a reduced tobacco consumption trend that cannot be modelled with any degree of reliability. This is the growth in regulation, at every level of government, and in a wide variety of directions. Since the report of the U.S. Surgeon General in 1964 a countless number of non-smoking and anti-smoking initiatives have been adopted in every developed economy. For example, restrictions upon workplace smoking have been introduced at different times and with differing degrees of vigor in Canada’s provinces. In addition, voluntary codes of behavior have been adopted by the manufacturing companies at various points - the manufacturers may have seen such codes as less undesirable than tougher legislation. Restrictions have been imposed by school boards upon use of tobacco in schools, by municipalities on use in restaurants and public places under local jurisdiction etc. Restrictions on marketing and advertising from the federal government have been many and varied4.

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4For example, at the federal level, 1971 saw the introduction of health warnings on packaging on a ‘voluntary’ basis by manufacturers. Tar and nicotine content labelling was introduced in the 1980s. The Tobacco Products Control Act was
The standard approach to dealing with interventions is either to test for structural breaks or to use dummy variables - particularly where ‘large’ measures such as national legislation or regulation are implemented. No doubt a dummy specification search on the importance of some of these events would yield significant breaks in the data. Yet such events inevitably take time to influence behaviors which are strongly habit driven. The smoking environment has become ever tighter and more regulated in the past three decades. At some points in time, workplace and public place restrictions have proceeded slowly; at other points, they have been more abrupt. But even where significant legislation is passed, its implementation may be gradual. For example, workplace restrictions, embodied in provincial legislation, may require monitoring before compliance is achieved. From an econometric perspective this process can be interpreted as a stochastic trend. Accordingly, our trend variable will be interpreted primarily as reflecting the education and regulatory policies of governments at all levels.

We indicated in the introduction that the changing composition of the population - towards one composed increasingly of immigrants of non-European stock, whose smoking habits are radically different from both European immigrants and native born Canadians, may be responsible for part of the decline in consumption. No reasonable data are available on the quarterly composition of the population. And even if such a series were available, it would bear such a high correlation with the trend we use to model regulation and education that the presence of both variables would be problematic. Our strategy therefore is to estimate our models with a stochastic time trend and to net out the effects of population change ex-post - we can get reasonably good estimates of the population composition for two (near) end points in the data.

Our prior on Canada’s legislative ‘long march’ is that the recorded decline in the use implies that the demand elasticity for tobacco should have fallen. Consider the following representation of consumers who introduced in 1988 and its provisions were implemented gradually, starting in 1989. In 1997 the Tobacco Act was adopted, containing limitations on advertising, marketing and promotion. On the provincial front, numerous pieces of legislation have been adopted, ranging in purpose from denying access to youth purchases to limitations on use in the workplace. Details of all of these programs are to be found on the web site of the Canadian Council for Tobacco Control (www.cctc.ca) and the National Clearing House on Tobacco and Health programs (www.nchth.ca). For a survey of legislative changes since the year 2000, see the ‘policy and legislation’ section of Health Canada’s “The National Strategy: Moving Forward. 2002 Progress Report on Tobacco Control”.

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make up the cigarette market.

Each consumer has a demand for cigarettes of a particular intensity, measured by the parameter \( \theta \). Suppose we order consumers in accordance with their intensity of preferences, and further assume that demands are monotonic, or non-crossing. This is a standard way of modelling consumers in the public-utility pricing literature (e.g. Brown and Sibley, 1986). Total quantity is then the sum of individual quantities

\[
Q = \int_{\theta_{\text{min}}}^{\theta_{\text{max}}} q(p(t), \theta(\phi)) g(\theta) d\theta
\]

where the limits of integration refer to the lowest and highest levels of demand intensity, \( g(\theta) \) is the density of consumers with intensity parameter \( \theta \), \( q \) is individual demand, \( p \) is the price and depends upon taxes \( t \). The variable \( \theta \) is subject to legislation, workplace smoking restrictions etc. In short it can be influenced by public policy and we denote this by the variable \( \phi \) whose level is chosen by government. The presence of this influence in the lower limit of integration implies that a marginal increase in the value of \( \phi \) will reduce the number of smokers. The application of Leibnitz’s rule to this aggregate, \( \frac{\partial Q}{\partial \phi} \), yields two terms: one indicating the degree to which all continuing smokers reduce their consumption, the other indicating the change at the extensive margin in the number of smokers

\[
\frac{\partial Q}{\partial \phi} = \int_{\theta_{\text{min}}}^{\theta_{\text{max}}} \left( \frac{\partial q}{\partial \theta} \frac{\partial \theta}{\partial \phi} g(\theta) d\theta - \frac{\partial q_{\text{min}}}{\partial \phi} g(\theta_{\text{min}}) \right).
\]

Those who quit will be those with the least consumer surplus prior to the policy initiative.

A simple graphical illustration is given in Figure 3, involving an economy in which there are two smokers - big (b) and small (s) - who differ in their preference intensity. The low-demand smoker quits after the policy initiative because that measure reduces his surplus sufficiently - less pleasure is derived as a result of being warned that smoking is bad. In contrast, the high-demand smoker still has sufficient surplus to continue smoking at a lower rate. If the low-demand smoker had a higher demand elasticity it is to be anticipated (though not guaranteed), that the aggregate demand elasticity will have fallen after his departure. The evidence from tobacco surveys indicates that the number of cigarettes per smoker has fallen very little in comparison with the number of cigarettes per person in the population in recent decades. This observation supports the proposition here.
4 Econometric Analysis

4.1 Modeling Framework

A convenient starting point for estimation is the model proposed by Becker, Grossman and Murphy (1994).

Suppose that a representative agent maximizes her lifetime utility

$$\sum_{t=1}^{\infty} \beta^{t-1} U(C_t, C_{t-1}, Y_t, e_t)$$

subject to $$\sum_{t=1}^{\infty} \beta^{t-1} (Y_t + P_t C_t) = A_0$$, where $$C_t$$ is the cigarette consumption at time $$t$$, $$Y_t$$ is the consumption of a composite commodity, $$e_t$$ captures the impact of unmeasured life-cycle variables on utility, $$P_t$$ is the price of cigarettes, $$A_0$$ is the present value of wealth and $$\beta$$ is the discount rate. This optimization problem gives rise to the following first-order conditions with respect to $$C_t$$ and $$Y_t$$, respectively,

$$U_C(C_t, C_{t-1}, Y_t, e_t) + \beta U_C(C_{t+1}, C_t, Y_{t+1}, e_{t+1}) = \lambda P_t$$

$$U_Y(C_t, C_{t-1}, Y_t, e_t) = \lambda,$$

where $$\lambda$$ is the marginal utility of wealth. If the utility function is assumed to be quadratic in $$C_t, Y_t$$ and $$e_t$$ (Becker, Grossman and Murphy, 1994), solving the first-order conditions for $$C_t$$ produces a linear equation

$$C_t = \theta C_{t-1} + \beta \theta C_{t+1} + \theta_1 P_t + \theta_2 e_t + \theta_3 e_{t+1}, \quad (1)$$

where the parameters $$\theta, \theta_1, \theta_2$$ and $$\theta_3$$ are functions of $$\lambda, \beta$$ and the coefficients of the quadratic utility function.

The habit formation or the addictive behavior of cigarette consumers is captured by the parameter $$\theta$$. If $$\theta > 0$$, an alternative approach to modeling habit formation with more flexible utility specification can be borrowed from the asset pricing literature in which a representative consumer maximizes

$$\sum_{t=1}^{\infty} \beta^{t-1} E \left[ \frac{(C_t + \theta C_{t-1})^{1-\alpha}}{1-\alpha} | I_t \right], \quad \alpha > 0,$$

where $$C_t$$ is consumption at time $$t$$, $$I_0$$ is the information set at time 0, $$\alpha$$ is a risk preference parameter and $$\theta$$ captures some time nonseparability in preferences (Abel, 1990; Constantinides, 1990; Heaton, 1995; among others). The Euler equation for the consumer’s problem is given by

$$E \left[ \beta \left( S_{t+1}^{-\alpha} + \beta \theta S_{t+1}^{-\alpha} \right) R_{t+1} - \left( S_t^{-\alpha} + \beta \theta S_t^{-\alpha} \right) | I_t \right] = 0$$

where $$S_t = C_t + \theta C_{t-1}$$ and $$R_t$$ is the risk-free rate of return. Then, the unknown parameters can be estimated using the generalized method of moments.
the cigarettes are addictive and the degree of addiction is greater when $\theta$ is larger.

The empirical specification of equation (1) has the following form

$$C_t = \alpha_0 + \alpha_1 C_{t-1} + \alpha_2 C_{t+1} + \alpha_3 P_t + \alpha_4 Y_t^d + u_t,$$

where $Y_t^d$ is the real per capita disposable income. At the risk of deviating from the original rational addiction model of Becker, Grossman and Murphy (1994), we specify equation (2) in log-levels\(^6\). The logarithmic transformation tends to stabilize the residual variance and has the attractive feature that the coefficients are directly interpretable as elasticities. A logarithmic specification of (2) is also estimated by Galbraith and Kaiserman (1997).

### 4.2 Estimation Procedure

#### 4.2.1 Specification in Levels

The direct estimation of model (2) may give rise to some misleading results and needs to be done with extreme caution\(^7\). First, the unmeasured life-cycle variables $e_t$ and $e_{t+1}$ in the composite error term $u_t$ affect consumption at all periods and $C_{t-1}$ and $C_{t+1}$ should therefore be treated as endogenous. A more serious problem arises from the possible nonstationarity of the regressors and presence of lead and lag levels of the dependent variable among them. It is well known that if the variables are nonstationary, the parameters on the lag values of the dependent variable are consistently estimable even in the presence of serial correlation in the errors. However, if cigarette consumption is integrated of order one and is cointegrated with real cigarette prices and disposable income, so are its lag and lead values. This implies that some linear combination of the regressors may be stationary and the remaining cigarette consumption variable would pick up part of the autocorrelation in the process. As a result, the estimates of the parameters on the past and future consumption that provide some information about rational addiction as well as the price and

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\(^6\)We also estimate the equation in the original level specification. The price and income elasticities are evaluated at the sample means of the data and the corresponding standard errors are computed using the delta method. The results are similar to those obtained for the log specification and are not reported.

\(^7\)Some Monte Carlo simulation results that point to several pitfalls in estimating rational addiction models are available from the authors upon request.
income elasticities may be highly misleading. In a slightly different setup, Auld and Grootendorst (2002) show that the standard rational addiction model leads to spurious evidence in favor of the rational addiction hypothesis.

Fortunately, model (2) can be reparameterized in a more convenient form. After adding $\alpha_1 C_t - \alpha_1 C_t$ and $\alpha_2 C_t - \alpha_2 C_t$ to the right-hand side of (2) and collecting the common terms yields

$$C_t = \gamma_0 + \gamma_1 \Delta C_t + \gamma_2 \Delta C_{t+1} + \gamma_3 P_t + \gamma_4 Y^d_t + \varepsilon_t,$$

(3)

where $\gamma_0 = \alpha_0/(1 - \alpha_1 - \alpha_2)$, $\gamma_1 = -\alpha_1/(1 - \alpha_1 - \alpha_2)$, $\gamma_2 = \alpha_2/(1 - \alpha_1 - \alpha_2)$, $\gamma_3 = \alpha_3/(1 - \alpha_1 - \alpha_2)$ and $\gamma_4 = \alpha_4/(1 - \alpha_1 - \alpha_2)$. This reparameterization allows us to separate the regressors into stationary and nonstationary components and the new parameters in front of $P_t$ and $Y^d_t$ have the interpretation of long-run elasticities. Park and Phillips (1989) show that the OLS estimators for the coefficients $\gamma_1$ and $\gamma_2$ are inconsistent if stationary regressors ($\Delta C_t$ and $\Delta C_{t+1}$) are correlated with the errors, whereas the estimators for the nonstationary coefficients are super- (rate $T$) consistent and their consistency is unaffected by the presence of endogenous stationary components.

Since our primary objective is to study the behavior of the long-run price elasticity of the demand, the changes in cigarette consumption have no first-order effect on our parameter of interest and can be omitted from the regression. In principle, if the parameters on past and future consumption are also of direct interest, we can employ a two-step procedure. As a first step, we estimate the level model ignoring the stationary components. Since these estimates are super-consistent, they can be treated as known in the second stage. Provided that the residuals from the first stage are stationary, the coefficients on the stationary components can be estimated using the generalized method of moments (GMM) estimator and these estimates can be used to back out the parameters of interest $\alpha_1$ and $\alpha_2$ in model (2).

First we describe the method for estimating the long-run demand for cigarettes model with no stationary components and testing for instability in the elasticity estimates. To introduce the main idea, suppose we have the system:

$$y_t^* = x_t^\prime \alpha + \varepsilon_{1t}$$

$$x_t^* = x_{t-1}^* + \varepsilon_{2t},$$
where $y_t^*$ and $x_t^*$ denote the detrended series, $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t})'$ are I(0) processes with a nondiagonal long-run covariance matrix $\Omega = \Lambda + \Gamma'$ with $\Lambda = \Gamma_0 + \Gamma$, $\Gamma_0 = E(\varepsilon_1 \varepsilon_t')$ and $\Gamma = \sum_{k=1}^{T-1} E(\varepsilon_t \varepsilon'_{t-k})$. It is assumed that the partial sum process $S_t = \sum_{j=1}^{t} \varepsilon_j$ satisfies the conditions for the multivariate invariance principle $T^{-1/2} S_{[Tr]} \Rightarrow W(r)$ for $r \in [0,1]$, where $[Tr]$ denotes the largest integer less than or equal to $Tr$, $W(r) = (W_1(r), W_2(r))'$, $W_2(r) = W_2(r) - \int_{0}^{1}(4 - 6s)W_2(s)ds - r \int_{0}^{1}(12s - 6)W_2(s)ds$, $W_1(r)$ and $W_2(r)$ are correlated Brownian motions, and $\Rightarrow$ signifies weak convergence as $T \to \infty$. Also, it is convenient to partition the long-run covariance matrix as $\Omega = \begin{pmatrix} \omega_{11} & \omega_{12} \\ \omega_{12}' & \Omega_{22} \end{pmatrix}$ and define $\omega_{1,2} = \omega_{11} - \omega_{12} \Omega_{22}^{-1} \omega_{12}'$.

The OLS estimation on the first equation of the system yields consistent but biased estimates of $\alpha$ due to misspecified dynamics and the OLS estimator converges to a nonnormal random variable $T(\hat{\alpha} - \alpha) \Rightarrow (\int_{0}^{1} W_2'(r)W_2'(r)dr)^{-1}(\int_{0}^{1} W_2'(r)dW_1(r) + \Lambda_{21})$. The endogeneity bias term $\Lambda_{21}$ arises from the correlation between the lagged values of $\Delta x_t$ and $\varepsilon_t$, $\Lambda_{21} = \sum_{k=0}^{\infty} E(\Delta x_{t-k} \varepsilon_t) = E(x_0^* \varepsilon_0)$. Furthermore, the limiting representation of the OLS estimator is nonpivotal and depends on nuisance parameters.

To eliminate the endogeneity bias and the dependence of the asymptotic distribution on nuisance parameters, Phillips and Hansen (1990) proposed the fully modified OLS estimator\(^8\) given by

$$\hat{\alpha}_{FM} = \left( \sum_{t=1}^{T} x_t^* x_t'^* \right)^{-1} \left( \sum_{t=1}^{T} x_t^* y_t^* - T \hat{\Lambda}_{21}^+ \right),$$

where $y_t^* = y_t^* - \hat{\omega}_{12} \hat{\Omega}_{22}^{-1} \Delta x_t^*$, $\hat{\Lambda}_{21}^+ = T^{-1} \sum_{k=0}^{\infty} \sum_{t=k+1}^{T} \Delta x_{t-k}^* \hat{\varepsilon}_t - \hat{\omega}_{12} \hat{\Omega}_{22}^{-1} \hat{\Lambda}_{22}$, $\hat{\varepsilon}_t$ denotes the OLS residuals and $l \to \infty$ as $T \to \infty$ such that $l = O(T^{1/4})$.

To compute the estimator $\hat{\alpha}_{FM}$, we need a consistent estimate of the long-run covariance matrix $\Gamma = \lim_{T \to \infty} T^{-1} \sum_{t=2}^{T} \sum_{j=1}^{t-1} E(\varepsilon_t \varepsilon'_{j})$. We follow Andrews and Monahan (1992) and use the ‘prewhitened’ resid-

\(^8\)In another method that corrects for the endogeneity bias, Stock and Watson (1993) orthogonalize the system by projecting $u_{1t}$ on $u_{2t}$ which gives $u_{1t} = E(u_{1t} | u_{2t}) + \varepsilon_t = d(L)u_{2t} + \varepsilon_t = d(L) \Delta x_t^* + \varepsilon_t$, where $d(L)$ is a two-sided (leads and lags) polynomial. Then the cointegrating vector $\alpha$ can be estimated efficiently from the dynamic OLS (DOLS) regression $y_t^* = x_t^* \alpha + d(L) \Delta x_t^* + \varepsilon_t$.

Both the DOLS of Stock and Watson (1993) and the fully modified estimator of Phillips and Hansen (1990) are single equation methods and depend on the chosen normalization of the cointegrating vector (Ng and Perron, 1997). Alternatively, we can use the full-information maximum likelihood method suggested by Johansen (1988).
uls $\hat{\xi}_t$ from a VAR(1) on the OLS residuals $\hat{e}_t$ and then compute

$$\hat{\Lambda}_\xi = \frac{1}{T} \sum_{t=1}^{T} \xi_t \hat{\xi}_t' + \frac{1}{T} \sum_{j=1}^{m} K(j/m) \sum_{t=j+1}^{T} \hat{\xi}_{t-j} \hat{\xi}_{t-j}'$$

$$\hat{\Omega}_\xi = \frac{1}{T} \sum_{t=1}^{T} \xi_t \hat{\xi}_t' + \frac{1}{T} \sum_{j=1}^{m} K(j/m) \left( \sum_{t=j+1}^{T} \hat{\xi}_{t-j} \hat{\xi}_{t-j}' + \hat{\xi}_{t-j} \hat{\xi}_{t-j}' \right),$$

where $K(j/m)$ is a weight (kernel) function and $m$ is the bandwidth parameter. In the empirical part of the paper, we use the quadratic spectral kernel and the plug-in bandwidth parameter recommended by Andrews (1991). Finally, we ‘recolor’ $\hat{\Lambda}_\xi$ and $\hat{\Omega}_\xi$ to obtain estimates of the covariance matrices of interest $\hat{\Lambda}$ and $\hat{\Omega}$.


Our interest also lies in testing whether the elasticity estimates in (3) are stable over time. For this purpose, we use several tests for parameter instability in regressions with integrated variables proposed by Hansen (1992). The first two statistics test the hypothesis of a single structural break in the coefficient vector $\alpha$ with the timing of the structural break being treated as unknown. The form of these statistics is

$$SupF_T = \sup_{t \in [L,T]} F_T(t)$$

$$AveF_T = \frac{1}{T-L+1} \sum_{t=L}^{T} F_T(t),$$

where $F_T(t) = \sum_{t=1}^{t} (x_t' \tilde{\epsilon}_t - \tilde{\Lambda}_{21}^{-1})$, $\tilde{\epsilon}_t = \hat{\epsilon}_t - \hat{\omega}_{12} \hat{\Omega}_{22}^{-1} \Delta x_t$, $V_T(t) = M_T(t) - M_T(t)M_T(T)^{-1}M_T(t)$, $M_T(t) = \sum_{t=1}^{T} x_t' x_t$, $L = \lfloor \pi T \rfloor$, $T = \lfloor (1-\pi)T \rfloor$ and $\pi \in (.1,.3)$.

The third statistic tests the hypothesis that the parameter vector $\alpha$ is changing smoothly over time and is given by

$$L = tr \left\{ M_T(T)^{-1} \sum_{t=1}^{T} S_T(t)' S_T(t) \right\} \hat{\Omega}_{12}^{-1}.$$
the infimum of the augmented Dickey-Fuller and Phillips-Perron tests over all possible regime shifts in the interval \([.15T, .85T]\). Then the values of the test statistics are compared to the asymptotic critical values in Gregory and Hansen (1996) that are derived under the null of no cointegration.

4.2.2 Specification in First Differences

The analysis in the previous section depends crucially on the assumption that the errors in \(2\) are stationary, i.e. the variables are cointegrated. If the cointegration is not supported by the data, the estimation of the level model would give rise to spurious results. More specifically, Lewbel and Ng (2002) argue that aggregating over a slowly evolving population with heterogeneous preferences might generate a highly persistent, nearly integrated error component. If the error term is a near unit root process, the hypothesis of no cointegration can not be rejected with probability approaching one and we need to induce stationarity in the variables to avoid spurious results.

To induce stationarity, we transform the model in first differences. This transformation also removes the jumps in the level of the price index for cigarettes that occurred in the beginning of 1994. To account for the possible endogeneity of the regressors as well as the presence of serial correlation and heteroskedasticity in the error term, we adopt the generalized method of moments (GMM) estimation framework.

Consider the first-differenced form of the rational addiction model. Let \(y_t = \Delta C_t\) and \(x_t = (1, \Delta C_{t-1}, \Delta C_{t+1}, \Delta P_t, \Delta Y_t^d)^\prime\) for the first-differenced form of \(2\) and \(x_t = (1, \Delta \Delta C_t, \Delta \Delta C_{t+1}, \Delta P_t, \Delta Y_t^d)^\prime\) for the first-difference form of \(3\). Suppose that, due to the endogeneity of (some of) the regressors, \(E(\varepsilon_t|x_t) \neq 0\) but we observe a set of instrumental variables \(z_t\) that satisfy the conditional moment restriction \(E(\varepsilon_t|z_t) = 0\).

The GMM estimator minimizes a quadratic form of the sample counterparts of the unconditional moment restrictions and is given by

\[
\hat{\theta} = \arg \min_{\theta \in \Theta} Q_T(\theta)
\]

\[
Q_T(\theta) = g_T(\theta)'W_Tg_T(\theta),
\]

where \(g_T(\theta) = \frac{1}{T} \sum_{t=1}^{T} z_t(y_t-x_t'\theta) = \frac{1}{T}(Z'Y-Z'X \theta)\) and \(W_T\) is a positive definite weighting matrix. Solving
the first-order conditions gives the GMM estimator

$$\hat{\theta} = (X'ZW_TZ'X)^{-1}X'ZW_TZ'Y. \quad (8)$$

The properties of the GMM estimator depend crucially on the choice of the weighting matrix. The efficient GMM estimator sets

$$W_T = \hat{V}_T^{-1},$$

where \(\hat{V}_T\) is a consistent estimate of the long-run variance of the moment conditions

$$V = \lim_{T \to \infty} E\left[\left(\frac{1}{\sqrt{T}} \sum_{t=1}^{T} z_t \epsilon_t \right) \left(\frac{1}{\sqrt{T}} \sum_{t=1}^{T} z_t \epsilon_t \right)'\right].$$

This variance-covariance matrix accounts for both heteroskedasticity and serial correlation of general unknown form in the moment conditions. By contrast, the familiar two-stage least square (2SLS) estimator sets

$$W_T = (Z'Z)^{-1}$$

and is inefficient in the presence of heteroskedasticity and autocorrelation.

In the estimation, we use the heteroskedasticity and autocorrelation consistent estimator

$$\hat{V}_T = \Gamma(0) + \sum_{j=1}^{m} K(j/m) \left(\hat{\Gamma}(j) + \hat{\Gamma}(j)'ight),$$

where \(K(j/m)\) is the quadratic spectral kernel, \(\hat{\Gamma}(j) = E[\eta_t \eta_{t-j}]\), \(\eta_t\) is obtained by applying a prewhitening filter to \(z_t \epsilon_t\), and \(m\) is the bandwidth which is determined using the automatic (data-dependent) bandwidth procedure of Andrews (1991). Under some regularity conditions, the efficient GMM estimator is consistent, \(\hat{\theta} \xrightarrow{p} \theta\), and asymptotically normal,

$$\sqrt{T}(\hat{\theta} - \theta) \xrightarrow{d} N\left(0, (M'V^{-1}M)^{-1}\right),$$

where \(M = \frac{\partial}{\partial \theta} Q(\theta)\).

### 5 Estimation Results

We begin the empirical analysis with determining the order of integration of cigarette consumption per capita, real price of cigarettes and real disposable income per capita. The deterministic component of each of the series is removed by GLS detrending (Elliott, Rothenberg and Stock, 1996) and the number of lags included in the augmented Dickey-Fuller regression is determined by the modified AIC of Ng and Perron (2001).

Table 3 reports the GLS detrended versions of the modified Phillips-Perron (\(MZ_\alpha\) and \(MZ_l\)) tests and the augmented Dickey-Fuller (ADF) test. The unit root tests show unambiguously that these series contain a unit root, i.e. they are I(1) processes. On the other hand, the first differences of cigarette consumption, price of cigarettes and real disposable income appear to be stationary or I(0) processes.

The first panel of Table 4 reports the estimates from the level specification of the rational addiction
model. The model includes a time trend and is estimated by the FM-OLS method of Phillips and Hansen (1990).

The negative estimate on the trend variable indicates a declining cigarette consumption of 0.8% per quarter after controlling for price and income effects. Several possible candidates to explain this nontrivial decline in cigarette consumption include anti-smoking advertising and regulation, tastes and demographic changes in Canadian population over time. As we pointed out in the introduction, the Canadian immigration patterns have changed radically since the sixties, away from European-source countries and towards immigrants from Asian, African and the Caribbean countries with significantly lower smoking rates. A more detailed data set will provide very valuable information in identifying the sources of this strong downward trend in cigarette consumption that we document here.

The sensitivity of the cigarette consumption to changes in the relative price of cigarettes and income deserves more careful analysis. The long-run price elasticity of cigarette consumption is -0.3 with a standard error of 0.029. This estimate suggests that smoking is somewhat insensitive to price changes which is consistent with the findings reported elsewhere (Galbraith and Kaiserman, 1997; Reinhardt and Giles, 2001; Gruber, Sen and Stabile, 2002; among others) although the previous studies report price elasticities between -0.4 and -0.5. The long-run income elasticity of cigarette consumption is 1.25 with a standard error of 0.125. This high income elasticity of demand is surprising, but nonetheless is not out of line with some estimates in the literature surveyed by Chaloupka and Warner (2000). High elasticities are obtained more frequently in time-series than in cross-section studies. Subsequently, we will see that income elasticity is time-varying and the reported high estimate in the sample is driven by a temporary increase in income sensitivity in the mid nineties.

The comparison of our findings to the previous results suggests that both price and income elasticities of demand for cigarettes may not be stable over time. Also, the second panel of Table 4 indicates that the data do not support the hypothesis of a constant cointegrating relationship between cigarette consumption, real price of cigarettes and real disposable income. Moreover, the parameter instability tests reject the hypothesis that the parameters of the model are stable over time\(^9\) which also confirms that the stable long-

\(^9\)The \(L\) test of a smoothly changing parameter vector does not reject the null at 5% but rejects at 10% significance level.
run cointegrating model is rejected by the data. For this reason, we computed the tests of no cointegration versus the alternative of cointegration with regime shifts. The results of the tests are reported in the bottom panel of Table 4. The Phillips-Perron tests strongly reject the null in favor of cointegration with possible structural break. The tests estimate the first quarter of 1991 as the time of the break. The results, however, are not fully conclusive since the ADF does not reject the null of no cointegration even though the value of the statistic is relatively close to its critical value.

To investigate further the stability of the cigarette sensitivity to price and income changes over time, we construct recursive and rolling sample estimates of the long-run price and income elasticities of cigarette consumption. In the recursive estimation the sample size is increasing whereas for the rolling sample estimation the sample size is held fixed with a rolling window of 50, 60, 70 and 80 observations, respectively. The first recursive estimate is constructed using the period 1972:Q1-1984:Q2 (50 observations), the second estimate is obtained from the period 1973:Q1-1984:Q3 (51 observations) etc. The rolling window estimation is expected to be more robust to potential structural shift that may have occurred due to legislative or tax policies. We also report the recursive and rolling estimates for the trend coefficient in the model.

The rolling sample and recursive sample results are plotted in Figures 4 to 9. Graphs 4 and 7 show that the sensitivity of cigarette consumption to price changes increased significantly in the eighties and steadily decreased in the nineties from -0.55 to -0.2 for the rolling sample estimates and to -0.3 for the recursive sample estimates. Figures 5 and 8 for the rolling and recursive estimates of income elasticity also offer some interesting results. Until the early nineties, the income elasticity is hovering around 0 being slightly positive in the eighties and slightly negative in the early nineties. The rolling sample estimates of size 70 and 80 indicate a sharp increase in income elasticity in mid nineties and subsequent decline in the second half of the decade. The recursive estimates do not capture the drop in the elasticity towards the end of the sample and show an increase of the income elasticity from 0 in 1994 to 1.25 by the end of the sample period. Finally, Figures 6 and 9 reveal that the strong downward trend in tobacco consumption, after the price and income effects are accounted for, has been formed around 1994 and stabilized at the value of -0.008 from 1996 onwards.
Since the results in Table 4 do not fully support the hypothesis of existence of a cointegrating relationship with regime shifts, there is a possibility that the results from the level specification are spurious. To check the robustness of the results, we use the first-differenced form of the model. This transformation also gives us some additional information about the impact and short-run elasticities of cigarette consumption. The first-differenced specifications of (2) and (3) are estimated by GMM. We follow Becker, Grossman and Murphy (1994) and use the future price change ($\Delta P_{t+1}$) as an instrument. The full set of instruments for both specifications is $(1, \Delta P_{t+1}, \Delta P_t, ..., \Delta P_{t-4}, \Delta Y^d_t, ..., \Delta Y^d_{t-4})$ which gives 7 overidentifying restrictions that can be used for specification testing. The first-differenced form of (3) appears to be more appealing since it gives directly the short-run elasticities and does not require further manipulations of the estimated coefficients.

The estimation results from the differenced specifications are given in Table 5. The short-run price elasticity estimated from these models (-0.11 and -0.13) suggests again that smoking is insensitive to price changes. The impact price elasticity is slightly higher (-0.29) and its magnitude is compatible with the long-run elasticity estimated from the level specification. Both impact and short-run income elasticities are not statistically significant and show that smokers do not adjust their tobacco consumption in response to changes in their income in the short-run. The $J$-test in both models can not reject the null that the models are correctly specified at 5% significance level. Note also that the estimates in front of the differenced quantity variable have no direct interpretation of rational addiction parameters. In fact, since the sum of the coefficients on $\triangle \triangle C_t$ and $\triangle \triangle C_{t+1}$ in specification I is not significantly different than zero, these two variables can be omitted and the model can be estimated by OLS.

In order to see if the short-run price elasticity exhibits a time-varying pattern that we observe for the long-run elasticity, we construct recursive and rolling sample estimates that are plotted in Figures 10 and 11. These graphs reveal that the behavior of the short-run elasticity over time is very similar to the time profile of the long-run elasticity. The short-run elasticity decreased significantly in the early nineties and its magnitude stabilized over the second half of the decade at values twice as low as the values in the eighties.
6 Conclusions and Public Policy

Our results provide several new insights on public policy. In the first place we find that, by the end of our sample period, the price elasticity of demand may be much lower than most previous estimates have indicated. Our recursive and rolling sample estimates yield values for the short- and long-run price elasticities in the range of \(-0.1, -0.3\). If we focus upon the years 2001/02, which are outside our sample period but during which prices increased by about 35%, then we should conclude that the aggregate elasticity is more likely in the less elastic portion of this region. Such a prediction is entirely consistent with the quantity data available at the time of writing: between the first half of 2000 and the first half of 2002 the huge price increase was accompanied by a quantity reduction of only about 7%.

This implies that price policies may have much more limited impact on aggregate demand than we previously thought. The tax revenue implications are equally stark.

As a public policy measure however, we need to look at the effectiveness of pricing policies on different demographic groups. There is ample evidence for the US that teen and youth smoking is more price sensitive than adult smoking: teens and youth have lower budgets, they experience stronger peer effects, and also are at a stage in the life cycle where they may not yet have become addicted. Accordingly, if youth rates have decreased by more that the aggregate figures indicate, a high-price policy may indeed be optimal for the long term, although it may take a generation for such a view to be verified.

A cautionary note however is necessary: when the market price of a good is driven strongly above its marginal cost of production by a tax wedge, substitutes can be brought to market readily. The experience of the early nineties stands as a clear case of this. At the time of writing this paper there is renewed evidence of contraband being produced and sold in the Akwesasne native Indian reserve that straddles the Ontario, Quebec and New York borders.

Second, we find evidence that regulation, education and other control measures have had a major impact on aggregate demand - over the long term perhaps by as much as three percent per annum. This is important because, while we have not sought to verify the impact of particular regulatory measures on consumption, it indicates that the struggle to reduce tobacco consumption may be a very long one. For example, the grim
health warnings on tobacco packaging that were introduced in 2001 were expected to have a major impact on purchases. But the micro data from Statistics Canada’s *Canadian Tobacco Use Monitoring Survey* do not indicate that the measure’s short term aggregate impact was at all significant. What the findings in this paper indicate is that the impact of control measures is slow, but nonetheless cumulative and decisive: public policy developers should not be surprised or perturbed that measures they hoped would constitute the ‘next big thing’ against consumption have but a modest effect in the short to medium term.

Third, we indicated in the introduction that some part of the behavior in the trend variable may be attributable to changing demographic composition. A recent report from Health Canada indicates that 16% of the Canadian population is currently of non-European immigrant stock, in contrast to about 5% in 1970. Given that this immigrant group has a smoking prevalence rate in the region of 10%, its growth over the time period can explain about three percentage points of the total drop in prevalence over the period - as much as one quarter of the total reduction. To our knowledge, this pattern has escaped the attention of analysts.

Fourth, we have emphasized the importance of using a meaningful definition of cigarettes - one that includes the hand-rolled substitutes for manufactured products. The lower degree of price elasticity that we find relative to other studies may be attributable in part to our definition of the quantity variable: price increases indeed cause some individuals to quit, but induce others merely to switch to a lower priced tobacco product. Omitting this component can lead to the erroneous conclusion that more people quit in response to high-tax policies than actually do.

For policy purposes the limitations of the findings should be apparent. The results provide no support for any specific measures enacted with the purpose of reducing consumption. They provide support for all of the measures as a set. However, the results do support the stance that tax increases may be very lucrative for governments. What remains surprising is the limited degree to which the manufacturers have increased price in this oligopolistic market structure - an issue that has been investigated by Harris (1983), Porter (1986) and Sumner (1981), but this issue is beyond the scope of our current focus.
References


Table 1. Manufactured and fine-cut tobacco sales, Canada 1999.

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<tr>
<th>Province</th>
<th>$C_m$(millions)</th>
<th>$C_{fc}$(kg)</th>
<th>$\frac{C_{fc}}{C_m+C_{fc}}$</th>
<th>$C_m$ taxes/200</th>
<th>$C_{fc}$ taxes/200gr</th>
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Table 2. Manufactured and fine-cut tobacco sales, Canada 1986

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<td>5.09</td>
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<tr>
<td>NB</td>
<td>1,361.8</td>
<td>448,700</td>
<td>0.276</td>
<td>14.13</td>
<td>5.18</td>
</tr>
<tr>
<td>Que</td>
<td>14,248.8</td>
<td>2,773,300</td>
<td>0.184</td>
<td>15.17</td>
<td>5.62</td>
</tr>
<tr>
<td>Ont</td>
<td>20,489.7</td>
<td>1,418,400</td>
<td>0.074</td>
<td>11.53</td>
<td>4.62</td>
</tr>
<tr>
<td>Man</td>
<td>2,851.5</td>
<td>369,500</td>
<td>0.130</td>
<td>14.33</td>
<td>4.82</td>
</tr>
<tr>
<td>Sask</td>
<td>1,797.4</td>
<td>265,700</td>
<td>0.146</td>
<td>14.29</td>
<td>4.92</td>
</tr>
<tr>
<td>Alta</td>
<td>6,351.0</td>
<td>410,600</td>
<td>0.070</td>
<td>9.09</td>
<td>2.90</td>
</tr>
<tr>
<td>BC</td>
<td>5,435.9</td>
<td>777,500</td>
<td>0.142</td>
<td>13.01</td>
<td>4.58</td>
</tr>
<tr>
<td>Total</td>
<td>55,435.9</td>
<td>7,353,500</td>
<td>0.133</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
Table 3. Unit root tests

<table>
<thead>
<tr>
<th></th>
<th>C</th>
<th>ΔC</th>
<th>P</th>
<th>ΔP</th>
<th>Y^d</th>
<th>ΔY^d</th>
</tr>
</thead>
<tbody>
<tr>
<td>$MZ_{\alpha}$</td>
<td>-1.14</td>
<td>-9.27</td>
<td>-3.46</td>
<td>-13.12</td>
<td>-2.21</td>
<td>-12.87</td>
</tr>
<tr>
<td>$MZ_t$</td>
<td>-0.53</td>
<td>-2.10</td>
<td>-1.29</td>
<td>-2.56</td>
<td>-0.97</td>
<td>-2.45</td>
</tr>
<tr>
<td>$ADF$</td>
<td>-0.92</td>
<td>-4.31</td>
<td>-1.32</td>
<td>-3.09</td>
<td>-1.01</td>
<td>-4.27</td>
</tr>
</tbody>
</table>

Note: The levels of all series are GLS detrended and the first differences of all the series are GLS demeaned. The 5% critical values for the unit root tests in levels are -6.3, -2.86 and -2.86 for $MZ_{\alpha}$, $MZ_t$ and $ADF$, respectively. The corresponding 5% critical values for the tests in first differences are -8.1, -1.95 and -1.95.
Table 4. Estimation and test results from the level specification

<table>
<thead>
<tr>
<th></th>
<th>$P$</th>
<th>$Y^d$</th>
<th>trend</th>
</tr>
</thead>
<tbody>
<tr>
<td>estimates</td>
<td>-0.303</td>
<td>1.248</td>
<td>-0.0084</td>
</tr>
<tr>
<td>standard errors</td>
<td>0.029</td>
<td>0.125</td>
<td>0.0005</td>
</tr>
</tbody>
</table>

residual-based tests for cointegration

<table>
<thead>
<tr>
<th></th>
<th>$MZ_{\alpha}$</th>
<th>$MZ_t$</th>
<th>$ADF$</th>
</tr>
</thead>
<tbody>
<tr>
<td>test value</td>
<td>-7.82</td>
<td>-1.50</td>
<td>-2.54</td>
</tr>
<tr>
<td>5% CV</td>
<td>-25.72</td>
<td>-4.16</td>
<td>-4.16</td>
</tr>
</tbody>
</table>

parameter instability tests

<table>
<thead>
<tr>
<th></th>
<th>$sup F^F_T$</th>
<th>$AveF^F_T$</th>
<th>$L$</th>
</tr>
</thead>
<tbody>
<tr>
<td>test value</td>
<td>19.66</td>
<td>12.53</td>
<td>0.722</td>
</tr>
<tr>
<td>5% CV</td>
<td>17.3</td>
<td>7.69</td>
<td>0.778</td>
</tr>
</tbody>
</table>

tests for cointegration with regime shifts

<table>
<thead>
<tr>
<th></th>
<th>$inf Z_{\alpha}$</th>
<th>$inf Z_t$</th>
<th>$inf ADF$</th>
</tr>
</thead>
<tbody>
<tr>
<td>test value</td>
<td>-117.93</td>
<td>-10.60</td>
<td>-4.79</td>
</tr>
<tr>
<td>5% CV</td>
<td>-68.94</td>
<td>-6.00</td>
<td>-6.00</td>
</tr>
</tbody>
</table>
Table 5. GMM estimation for the first-differenced specifications

<table>
<thead>
<tr>
<th>Specification</th>
<th>$\Delta \Delta C_t$</th>
<th>$\Delta \Delta C_{t+1}$</th>
<th>$\Delta P_t$</th>
<th>$\Delta Y_t^d$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Specification I</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Estimates</td>
<td>0.316</td>
<td>-0.330</td>
<td>-0.110</td>
<td>-0.025</td>
</tr>
<tr>
<td>Standard errors</td>
<td>0.023</td>
<td>0.034</td>
<td>0.010</td>
<td>0.089</td>
</tr>
<tr>
<td>$p$-value of $J$-test</td>
<td>= 0.072</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

| Specification II |                      |                          |             |               |
| Estimates        | -0.731               | -0.474                   | -0.291      | 0.335         |
| Standard errors  | 0.073                | 0.149                    | 0.025       | 0.275         |
| $p$-value of $J$-test | = 0.092          |                          |             |               |

Note: The instruments used in both specifications are $(1, \Delta P_{t+1}, \Delta P_t, ..., \Delta P_{t-4}, \Delta Y_t^d, ..., \Delta Y_{t-4}^d)$. 


Figure 1: Total cigarette sales per capita (seasonally adjusted).
Figure 2: Real price of cigarettes.
Figure 3: Demand reducing policies and aggregate demand elasticity.

*Note*: Each consumer displays a diminished demand as a result government health messages. The demands of the small ($s$) and big ($b$) consumers decline from $d_s$ and $d_b$ to $d_s^1$ and $d_b^1$ respectively. But at the (constant) price $P_0$ the small consumer no longer obtains any consumer surplus and therefore exits the market. The new aggregate demand is therefore simply the demand of the big consumer, and may be less elastic due the departure of the more price-elastic consumer.
Figure 4: Rolling sample estimates of the long-run price elasticity of demand for cigarettes.
Figure 5: Rolling sample estimates of the long-run income elasticity of demand for cigarettes.
Figure 6: Rolling sample estimates of the trend of demand for cigarettes.
Figure 7: Recursive sample estimates of the long-run price elasticity of demand for cigarettes.
Figure 8: Recursive sample estimates of the long-run income elasticity of demand for cigarettes.
Figure 9: Recursive sample estimates of the trend of demand for cigarettes.
Figure 10: Recursive sample estimates of the short-run price elasticity of demand for cigarettes.
Figure 11: Rolling sample estimates of the short-run price elasticity of demand for cigarettes.