Are Cigarette Bans Really Good Economic Policy?

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Abstract:

We investigate the quarterly relationship between the quantity of cigarettes sold, real disposable income *per capita*, and the relative price level of cigarettes in Canada. Careful attention is paid to the non-stationarity of the data and the dynamic specification of the model. We conclude that cigarette demand is extremely insensitive to price and income changes. This is evidence of the large consumer surplus smokers enjoy and the large revenue increasing potential of a cigarette tax increase policy, as opposed to cigarette bans.

Keywords: **Cigarettes; Smoking; Demand Function**

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I. INTRODUCTION

Twenty-five years ago, a significant public knowledge of the costs that smokers bestow on the rest of society, from the act of smoking, did not exist as it does today. Smoking affects the length and quality of life of those who do and do not smoke. Many medical dollars are devoted to the treatment of smoking-related sicknesses in Canada, and other countries, each year. Consequently, many anti-smoking policies have emerged[1], most of which are imposed with little or no regard for the large consumer surplus smokers enjoy form their cigarette consumption. The conversions of Toronto and Victoria to non-smoking cities are strong cases in point. The act of smoking has costs and benefits. The high visibility of the costs of smoking makes it politically tempting to ignore the value of cigarettes to smokers altogether when devising anti-smoking policies. However, smokers derive utility from the act of smoking (even if they smoke due to addiction), and their welfare matters.

In this paper we estimate a Canadian demand function for cigarettes and use it to demonstrate the consequence of a tax increase as an alternative policy to cigarette bans; the latter simply dictating how much and where to smoke, independent of differing marginal valuations.

II. DATA ISSUES

Our three quarterly time series variables are millions of cigarettes sold (Qd), real disposable income *per capita* (Yd), and the *relative* price level of cigarettes (RP). Each series are from 1968.1 to 1990.2 inclusive, the longest period available, and are obtained from the "CANSIM" database, Statistics Canada (1990, 1997, 1998a, 1998b, 1998c).

A notable feature of our cigarette sales and relative price variables is the structural break that occurs at around 1981.3. The relative price level of cigarettes grows quickly from that point forward and cigarette sales shrink. This structural break needs to be taken into account in various parts of our analysis.

First, we test for the presence of unit roots in the logs of all three variables at the 0, π, and $\pi/2$ frequencies using the Hylleberg, Engle, Granger and Yoo (1990) (HEGY) testing framework. In the case of annual data, if we ignore a structural break, the "augmented" Dickey-Fuller test would be biased toward concluding non-stationarity. One can deal with this by using Perron's (1989) modified asymptotic test. Unfortunately, there is no corresponding test available for stochastic seasonality, so we bootstrap the critical values of our HEGY tests to allow for the structural break and particular sample size.

For simplicity, we assume that each logged variable is integrated of order at most two (I(2)), and so we apply the HEGY testing framework to both the differences and levels of the (log-) data. A trend and three seasonal dummy variables are included in each HEGY regression, and the numbers of augmentation terms are chosen to remove any significant autocorrelations and partial autocorrelations in the HEGY residuals. The null hypothesis of non-stationarity (existence of unit roots) at each individual frequency is tested with HEGY-type "t-tests". To further support our conclusions we also use joint "F-tests" for non-stationarity at the $\pi/2$ frequency, at all seasonal frequencies, and at all frequencies (*e.g.*, Hylleberg *et al*. (1990) and Ghysels, Lee and Noh (1994)). The critical values are bootstrapped using 2,000 repetitions under the "S_2" sampling scheme suggested by Li and Maddala (1996).

The results of our HEGY testing appear in Tables 1 and 2. We find that log(Qd), log(RP) and log(Yd) are integrated of order 1, 1 and 2 respectively at the 0 frequency. Log(Yd) is also integrated of order 1 at the $\pi/2$ frequency. This suggests that there is the potential for (first-order) cointegration only at the 0 frequency between all three variables.

III. TESTING FOR COINTEGRATION

We follow the procedure outlined by Engle, Granger, Hylleberg and Lee (1993) (EGHL). The series for log(Qd) and log(RP), filtered for the non-zero frequencies, from our HEGY testing are used again here. As log(Yd) is apparently integrated of order two, we need to first difference the series and then filter it for the non-zero frequencies in the usual way. The three filtered series are as follows:

$$Q_{1t}=(1+L)(1+L^2)\Delta_4\log(Qd)$$

$P_{1t} = (1+L)(1+L^2)\Delta_4 \log(RP)$

$Y_{1t} = (1+L)(1+L^2)\Delta_4[\Delta \log(Yd)]$,

where "L", "Δ" and "Δ_4" are the lag, first-differencing, and fourth-differencing operators.

Our first cointegrating regression, with the inclusion of three seasonal dummy variables (S_1, S_2 and S_3) and a trend variable (t), is:

$$\mathbf{Q_{1t} = a + b_1 P_{1t} + b_2 Y_{1t} + f\,t + d_1 S_{1t} + d_2 S_{2t} + d_3 S_{3t} + e_t .} \qquad \mathbf{(1)}$$

We obtain the OLS residuals, e_t, and then fit the auxiliary regression:

$$\mathbf{De_t = g e_{t-1} + S\ d_i De_{t-i} + v_t .} \qquad \mathbf{(2)}$$

The number of lagged dependent variables in (2) is chosen to ensure that there are no significant autocorrelations or partial autocorrelations for the residuals of this regression. We test $H_o: \gamma=0$ (no cointegration), against $H_A: \gamma<0$. The resulting t-statistic is –2.352, and the exact 10% critical value from MacKinnon (1991) is –3.946, so we conclude that no cointegration exists. This same procedure is applied with each variable as the dependent variable in (1) without altering our conclusion. We also test for the existence of bivariate cointegration between log(Qd) and log(RP). Our final conclusion is that cointegration does *not* exist, so an error-correction model is not appropriate here.

IV. The Model and Empirical Results

In the absence of cointegration we filter each variable to make it stationary:

$\Delta Q_t = (1-L)[\log(Qd_t)]$; $\Delta P_t = (1-L)[\log(RP_t)]$; and $Y^*_t = (1-L)^2(1+L^2)[\log(Yd_t)]$.

After fitting a relatively general dynamic model, we "tested down" to obtain the following final fitted OLS regression:

$$\mathbf{DQ_t = 0.118 - 0.617\ DP_t + 0.193\ Y^*_t - 0.038\ DUM_t - 0.223\ S1_t - 0.132\ S2_t - 0.031\ S3_t}$$
$$\mathbf{(0.02)\quad (0.22)\qquad (0.17)\qquad (0.01)\qquad (0.03)\qquad (0.04)\qquad (0.03)}$$

$$\mathbf{- 1.116\ DQ_{t-1} - 0.760\ DQ_{t-2} - 0.349\ DQ_{t-3} - 0.169\ DQ_{t-4} - 0.275\ DP_{t-1} - 0.274\ DP_{t-2}}$$
$$\mathbf{(0.12)\qquad (0.17)\qquad (0.15)\qquad (0.10)\qquad (0.23)\qquad (0.23)}$$

$$\mathbf{+ 0.118\ Y^*_{t-1}} \qquad \mathbf{(3)}$$
$$\mathbf{(0.16)}$$

Adjusted R^2 = 0.924; LM1=0.494; LM2=0.666; LM3=0.366; LM4=1.023

The numbers in parentheses are standard errors; DUM is a dummy variable which is zero before the structural break in 1982.3 and unity otherwise; and LMi is the LM statistic for testing serial independence against a simple ith-order autoregressive or moving average process in the disturbances. There is no significant evidence of autocorrelation and the satisfactory within-sample "fit" of the model is shown in Figure 1.

The coefficient on the (*differenced*) dependent variable lagged one period exceeds unity, but the *log-level* version of this fourth-order difference-equation is stable. First, (3) is re-written in log-level form. Then, following Hamilton (1994, pp. 7-16), we define the (4×4) matrix F, with rows $F_1 = (\phi_1, \phi_2, \phi_3, \phi_4)$; $F_2 = (1, 0, 0, 0)$; $F_3 = (0, 1, 0, 0)$ and $F_4 = (0, 0, 1, 0)$; where $\phi_1 = [1-(-1.116)] = -0.116$; $\phi_2 = [-0.7610-(-1.116)] = 0.355$; $\phi_3 = [-0.349-(-0.7610)] = 0.412$; and $\phi_4 = [-0.169-(-0.349)] = 0.180$. The eigenvalues of F are distinct, two of them being real (0.9415 and –0.5310) and the others being the complex conjugate pair, [-0.2632 +/- 0.5392 i], with modulus 0.6. As this latter value and the two real roots are each less than unity in absolute value, equation (3) is stable and non-oscillatory.

V. Economic Implications

The *short-run* price and income elasticity estimates and are η_{srp} = - 0.617 and η_{sry} = 0.193, respectively. These values imply the demand for cigarettes is price and income inelastic, according with the notion that necessity goods with few close substitutes have own-price inelastic demand curves.

Returning to the dynamics of model (3), as mod[-0.2632 +/- 0.5392 i] < 1, the multipliers associated with the estimated model result in impulses which are sinusoidal but with an amplitude which decays. The dynamic multipliers for price or income shocks can be calculated using formulae (1.2.24) and (1.2.25) of Hamilton (1994, p.12). These multipliers, or numerical simulation, can be used to examine the effect of a one period, and permanent, increase in price (for example) on the model's dynamic predictions of the demand for cigarettes. Such simulations mimic policies that increase cigarette taxes.

Let RP1 be the sample relative price vector under the current tax policy. By way of illustration, let RP2 equal RP1, except in 1974.1 where the price is artificially increased by 80% for that one period. Finally, let RP3 equal RP1 prior to1974.1, but be 80% higher than RP1 for the rest of the sample. The model's dynamic predictions of cigarette demand under these three hypothetical policy regimes are then derived by undifferencing and taking the anti-log of the predicted demand series. As can be seen in Figures 2 and 3, under both "policy on" scenarios, five quarters are needed for most of the adjustment to take place and four years are needed for full adjustment. After four years, cigarette demand under the policy invoking RP2 returns fully to the "policy off" case, associated with prices RP1. Under a policy invoking RP3, cigarette demand settles at 18.3% less than the "policy off" case in the long run (Figure 3). That is, the long-run price elasticity of demand is approximately $\eta_{lrp} = -0.228$.

VI. Conclusions

The presence of the external costs associated with smoking necessitates the need for policies designed to induce smokers to internalize the costs of their habit. If economic efficiency is the policy objective, it is essential that such policies do not ignore the value that smokers place on their cigarette consumption. Some smokers may even prefer death then to live without cigarettes. Our simulation experiments suggests that an 80% permanent increase in the price of cigarettes would only reduce the equilibrium level of cigarette consumption by a mere 18%. The implication of this is that smokers presently enjoy a large consumer surplus. By imposing a steep tax increase, cigarette consumption would be restricted to those who value it most. Further, the increase in tax revenue could be used to fully cover smoking health related costs and provide services, which are valued by both smokers and non-smokers.

Non-economic policies aimed at reducing smoking via cigarette bans or quotas ignore the fact that cigarettes are a highly valued commodity. Some may even object to these policies on the grounds of freedom of choice. As an alternative to dictating how much and where to smoke, the imposition of an increased sales tax on cigarettes can be used. This alternative can achieve the same aggregate reduction in cigarette consumption while still allowing individual agents to make choices at the margin. Under this policy option a portion of smokers surplus is converted into dollars which can then be funneled back into the economy via increased government expenditure. This is not possible under present legislation where smokers are simply forbidden to smoke in certain areas.

Table 1 – Seasonal Unit Root Tests, I(2) *vs.* I(1)

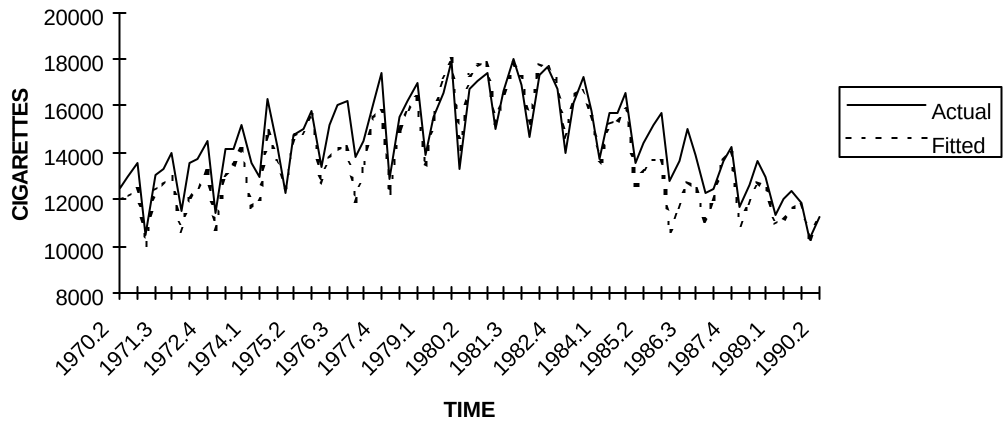
	Δlog(Qd)		Δlog(RP)		Δlog(Yd)	
Test Statistic	Actual Value	Bootstrapped 10% Critical Value(s)	Actual Value	Bootstrapped 10% Critical Value(s)	Actual Value	Bootstrapped 10% Critical Value(s)
$\mathbf{F_{1234}}$	43.459	4.67	11.86	4.18	6.576	4.01
$\mathbf{F_{234}}$	55.75	5.75	11.37	5.04	8.698	4.7
$\mathbf{F_{34}}$	18.75	4.44	10.02	4.36	5.146	4.64
$\mathbf{t_{p\,1}}$	-2.711	-1.67	-3.07	-1.67	-1.717	-2.08
$\mathbf{t_{p\,2}}$	-4.877	-2.69	-3.207	-2.62	-2.587	-2.51
$\mathbf{t_{p\,3}}$	-6.679	-3.13	-3.422	-2.88	-3.243	-2.45
$\mathbf{t_{p\,4}}$	2.328	-1.58, 1.17	-2.216	-1.93, 1.04	-0.666	-1.81, 1.89

Table 2 – Seasonal Unit Root Tests, I(1) *vs.* I(0)

	log(Qd)		log(RP)		Log(Yd)	
Test Statistic	Actual Value	Bootstrapped 10% Critical Value(s)	Actual Value	Bootstrapped 10% Critical Value(s)	Actual Value	Bootstrapped 10% Critical Value(s)
$\mathbf{F_{1234}}$	7.045	2.9	11.389	3.21	3.394	4.69
$\mathbf{F_{234}}$	8.397	3.16	10.138	3.6	4.381	5.54
$\mathbf{F_{34}}$	4.48	3.69	9.834	2.23	5.658	7.66
$\mathbf{t_{p\,1}}$	1.018	-2.35	-0.621	-2.23	-1.51	-3.4
$\mathbf{t_{p\,2}}$	-2.757	-1.57	-4.404	-1.43	-3.014	-2.13
$\mathbf{t_{p\,3}}$	-3.948	-1.62	-2.492	-2.5	-1.369	-1.54
$\mathbf{t_{p\,4}}$	1.384	-1.82, 1.75	-2.369	-1.92, 1.85	-0.395	-1.15, 2.05

Figure 1

Actual & Fitted Demand



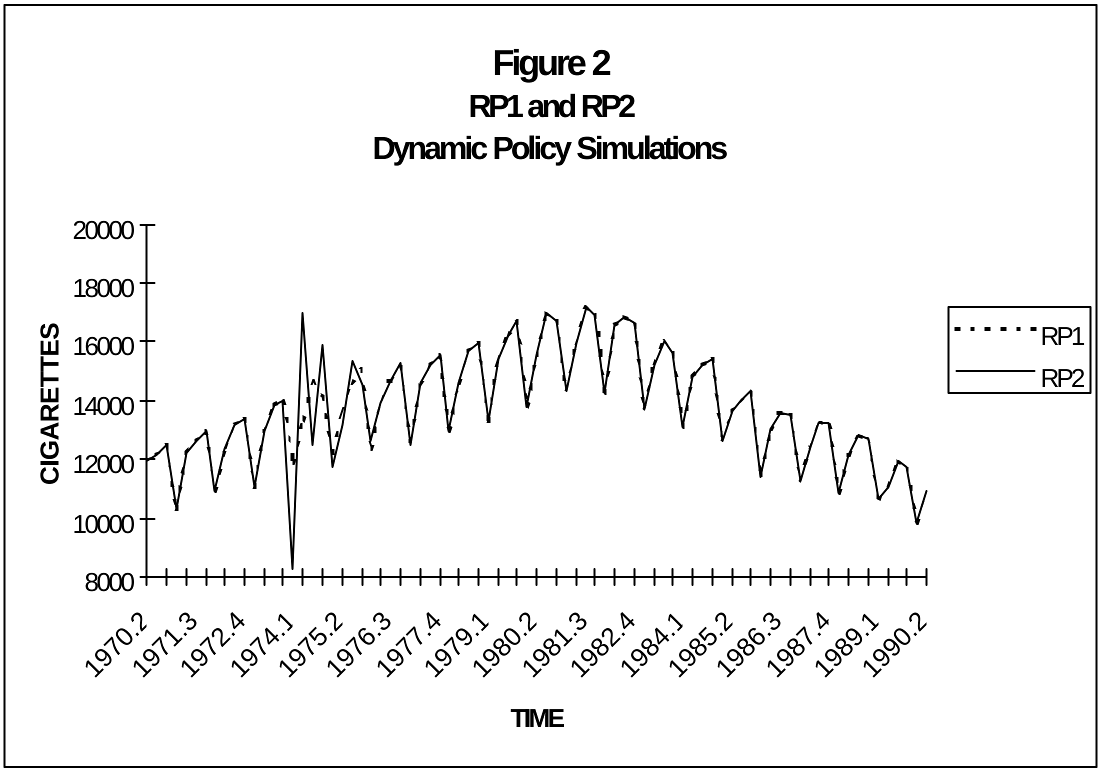


Figure 2
RP1 and RP2
Dynamic Policy Simulations
20000
18000
16000
14000
12000
10000
8000
CIGARETTES
1970.2
1971.3
1972.4
1974.1
1975.2
1976.3
1977.4
1979.1
1980.2
1981.3
1982.4
1984.1
1985.2
1986.3
1987.4
1989.1
1990.2
TIME
RP1
RP2

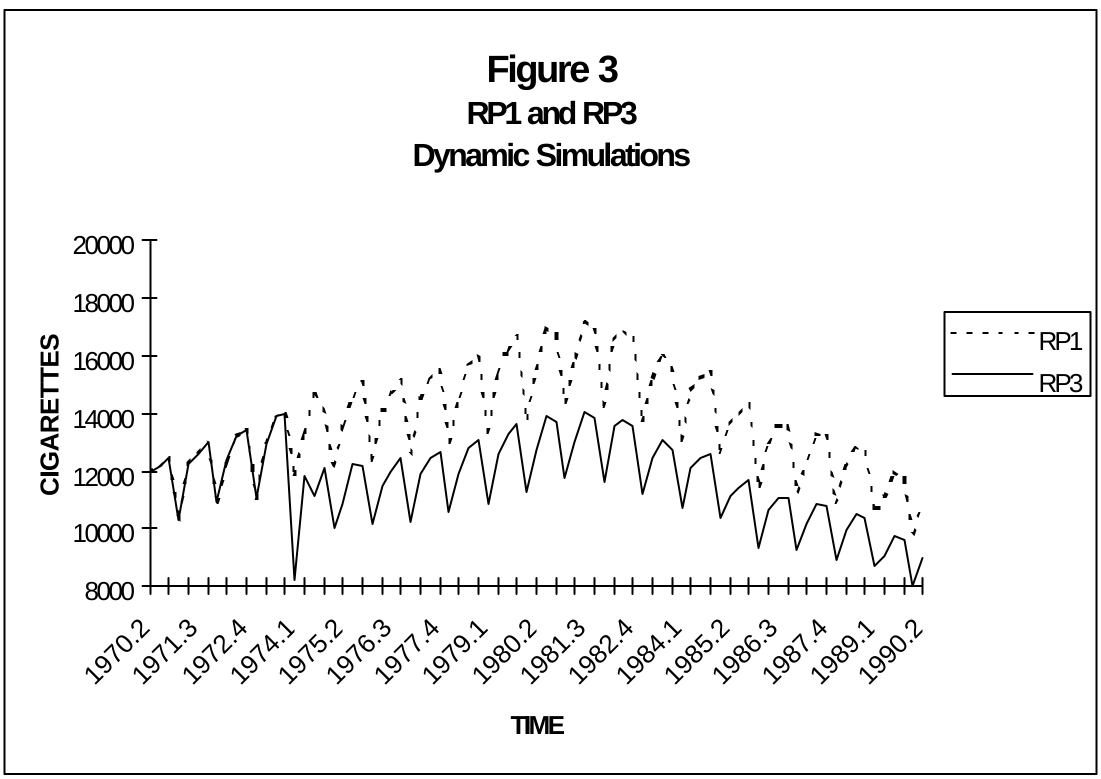


Figure 3
RP1 and RP3
Dynamic Simulations
20000
18000
16000
14000
12000
10000
8000
CIGARETTES
1970.2
1971.3
1972.4
1974.1
1975.2
1976.3
1977.4
1979.1
1980.2
1981.3
1982.4
1984.1
1985.2
1986.3
1987.4
1989.1
1990.2
TIME
RP1
RP3

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Footnote

1. See the url: www.tobaccofacts.org/. For example, in 1989, regulations requiring the new health warnings on cigarette packages came into effect.