New Evidence on Demand for Cigarettes:
A Panel Data Approach

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Abstract. This paper estimates the demand for cigarettes using panel data – 42 states and Washington, D.C. – from 1961 to 2002. We first employ the panel unit root test before estimating the demand structure. We have found that (i) the price and income elasticities are approximately –0.41 and 0.06, (ii) the price elasticities of neighboring states is 0.09, (iii) decreasing tax elasticity gives rise to decreasing price elasticity, and smaller tax shares (real tax as percentage of real price) seem to be related to declining tax elasticity, (iv) overall antismoking campaigns have contributed to declining income elasticities across different income groups, and (v) the decline in income elasticity for dividend and transfer income recipients is the main cause for the decrease in overall income elasticity. It has interesting implications: cigarette consumption is a normal good to wage earners and transfer payment recipients, but an inferior good to the owners of stocks and the elderly population.

Keywords: Cigarette demand, demand elasticity, panel data.

JEL classification: L66, C33

Introduction

Cigarette demand estimation has been of paramount interest since the Surgeon General's warning (1964) on the potential causal relationship between cigarette smoking and smoking related diseases, lung cancer in particular. Prior studies focused on the estimates of price, income, and advertising elasticities using either time series (e.g., Fujii, 1980; Bishop and Yoo, 1985; Keeler *et al.* 1993) or cross-section (e.g., Lyon and Simon, 1968) data. As pointed out by Baltagi and Levin (1986) in their cigarette demand study, the pure cross-section model cannot effectively control for state specific (demographic) effects whereas the pure time-series analysis fails to effectively control for unobservable taste changes. However, an appropriate pooled technique can provide opportunities to accommodate for both shortcomings.

Baltagi and Levin (1986) pioneered panel data estimation by incorporating 47 states over the period between 1963 and 1980. Baltagi *et al*. (2000) most recently discussed the homogeneous and heterogeneous estimator problems in the cigarette demand estimation using panel data. They strongly endorse pooled models, which posit a set of homogeneous parameters over their heterogeneous counterparts based on the root mean square error criterion. In their panel estimation (1960-1990), Keeler *et al*. (2001) emphasized the biased elasticity estimates emanating from omitted variables in demand functions. Nelson (2003) investigated the effects of advertising bans on cigarette consumption.

The panel data approach undoubtedly provides heterogeneous information at the state level while ensuring more reliable estimates via large sample property. With the growing popularity of panel data applications, the time series properties such as unit root and cointegration have received increasingly more attention. If a unit root exists in the panel data, the result using the level (logarithmic) is suspect. To the best of our knowledge, analyses of demand for cigarettes have thus far ignored the panel unit root or panel cointegration properties with the exception of Nelson (2003) and Bask and Melkersson (2004). The purpose of this paper is to (1) estimate cigarette demand structure after applying the panel unit root technique; (2) disaggregate income into earning, dividend, transfer payment, and estimate respective income elasticities; (3) partition the cigarette price into pre-tax price and tax in demand estimation; and (4) estimate the time-varying price and income elasticities via using the rolling window approach in an attempt to offer explanations on decreasing elasticities.

It should be noted that different state cigarette taxes are a hotbed for "bootlegging" across states. Such a cross-border effect is significant and cannot be ignored (Coats, 1995). To this end, we include average neighboring state prices into the model to capture the substitution effect. Anti-smoking campaigns date back to 1965 (warning label), 1969 (advertisement bans on television and radio), 1972 (U.S. Surgeon General's warning about environmental hazards from tobacco smoking), 1979 (U. S. Surgeon General's report concluding nicotine was addictive), 1986 (release of U. S. Surgeon General's report on involuntary smoking), 1988 (smoking ban on all flights of two hours or less), 1990 (smoking ban on all flights), and 1992 (release of the U. S. Environmental Protection Agency report on environmental tobacco smoking). These events may well have contributed to the volatility of price and income elasticity of cigarette demand over time and as such, need to be considered in evaluating the price and income elasticities estimated from a pooling model.

Most recently, Sloan *et al*., (2002) found an own price elasticity of -0.15 and income elasticity of 0.15 applying a two-stage least squares model over 1990-1998. Nelson (2003) employed a nonstationary panel model to improve on the Saffer-Chaloupka (2000) result. He found the price elasticity had increased from -0.4 to -0.621 with decreased income elasticity after the regime shift in 1985. Interestingly enough, Gallet (2003) estimated demand functions for the 45 states along with the conjectural variation coefficients and the Lerner indexes. While the majority of estimated average demand elasticity is less than unity, some are quite elastic: -2.78 (Louisiana), -2.73 (Alabama), and -1.84 (Kentucky). The estimated conjectural variation coefficients indicate an anti-competitive way of advertising, echoing findings by Tremblay and Tremblay (1995), which is in rough agreement with the sizes of the Lerner index. Kim and Seldon (2004) found the price elasticity of cigarette demand in South Korea to be -0.276 and a nearly zero income elasticity, which is comparable to -0.33 in Greece (Hondroyiannis and Papapetrou, 1997) and the average value of 22 OECD countries (Steward, 1993).

This paper employs the panel unit root technique for 42 states and Washington, D.C. before estimating cigarette demand. The null hypothesis of a unit root is easily rejected and as a result, the demand structure can be formulated in double logarithmic form. The next section introduces the model. The third section provides the data and panel unit root test result. The fourth section is the discussion of the empirical results, and the last section gives concluding remarks.

Research Model

According to Laughhunn and Lyon (1971), Hamilton (1972), Doron (1979), Baltagi and Levin (1986), and Baltagi, Griffin, and Xiong (2000), cigarette demand is formulated as below:

$$Q_{it} = f (P_{it}, Y_{it,} Pn_{it,} Z_{it}) \tag{1}$$

where Q_{it} is per capita sales of cigarettes by population in state i at time t; P_{it} is the retail price per pack; Y_{it} denotes per capita personal income; Pn_{it} represents the average cigarette price of neighboring states; and Z_{it} represents other explanatory variables[1] (*e.g.,* proportion of population ages 65 and older). All the variables are deflated by the consumer price index (CPI).

Two major strands of cigarette demand models dominate the literature: the partial adjustment model and the rational addiction model. The partial adjustment model emphasizes the habit formation phenomenon (past) while the rational addiction model argues that both of the past and future consumptions exert positive impacts on current consumption as developed by Becker and Murphy (1988). Consider the partial adjustment, or habit persistence model of cigarette demand:

$$\ln Q_{it} - \ln Q_{it-1} = \rho (\ln Q'_{it} - \ln Q_{it-1}) + \mu_{it} \tag{2}$$

in which $\ln Q^*_{it}$ represents the expected or desired consumption level as shown below:

$$\ln Q'_{it} = \alpha' + \beta'_1 \ln P_{it} + \beta'_2 \ln Y_{it} + \beta'_3 \ln Pn_{it} + Z'_{it} \gamma' \tag{3}$$

Substituting (3) into (2) immediately yields

$$\ln Q_{it} = \alpha_1 + (1-\rho) \ln Q_{it-1} + \beta_1 \ln P_{it} + \beta_2 \ln Y_{it} + \beta_3 \ln Pn_{it} + Z'_{it} \gamma + \mu_{it} \tag{4}$$

Empirical estimates based on the partial adjustment model are generally oriented towards short run elasticities and speed of adjustment in the long run (Baltagi and Levin, 1986 and 1992; Baltagi *et al.*, 2000).

Alternatively, Becker and Murphy (1988) developed the rational addiction hypothesis assuming: (1) individuals make choices that span over several time periods and (2) addictive consumptions in different time periods are complementary. Consequently, past and future consumptions enter the demand structure under the rational addiction paradigm as:

$$\ln Q_{it} = \alpha'_{it} + \alpha_1 \ln Q_{it-1} + \alpha_2 \ln Q_{it+1} + \beta_1 \ln P_{it} + \beta_2 \ln Y_{it} + \beta_3 Pn_{it} + Z_{it} \gamma + \mu_{it} \tag{5}$$

The rational addiction hypothesis is found to be consistent with cigarette demand by Chaloupka (1991), Becker *et al.* (1994), Labeaga (1993), Walters and Sloan (1995), and Escario and Molina (2001), Nelson (2003), and Sloan *et al.* (2002). However, it is deduced from long-term disaggregate individual behavior. Employing aggregate panel data (state), we estimate the demand structure of (4) based on the partial adjustment model.[2] It is to be pointed out the double-log functional form is employed for convenience. While incorporating interacting explanatory variables in a multiplicative form, its estimated elasticities are restricted to be constant. For

example, flexible functional form – the Box-Cox transformation – was used by Chang and Hsing (1991) in estimating residential electricity with the transformation parameter of 0.23. In a separate study, however, Hsing and Chang (2000) found demand for real M2 to be linear with the transformation parameter of 1.01. The Box-Cox model can indeed yield more accurate elasticities for future policy implementation and remains an interesting future research topic.

Data Source and Panel Unit Root Result

The cigarette data used in this study are obtained from *The Tax Burden on Tobacco* published by the Tobacco Institute. Data for 42 states and Washington D.C. over the 1961–2002 sample period include per capita cigarette consumption measured by tax paid per capita sales in packs (Q_{it}), per capita personal income deflated by 1982 – 1984 dollars (Y_{it}), weighted average real price (cents) per pack of premium brands deflated in 1982 – 1984 dollars (P_{it}), and weighted average real tax (state and federal cigarette taxes) per pack of cigarettes deflated by 1982 – 1984 CPI (Tax_{it})[3].

To take the cross-border bootlegging effect into consideration, we include the cigarette price(s) of neighboring states into the demand equation. Baltagi and Levin (1986) used the minimum price of neighboring states to capture the cross-border effect whereas Keeler, *et al.* (1993) employed the average price of neighboring states as the proxy. Well-documented in time series studies, economic variables are fraught with the unit root problem as was first pointed out by Nelson and Plosser (1982).[4] A common pitfall is to apply a regression without checking the unit root property at level with the ensuing high R^2 and low Durbin-Watson statistic: the spurious regression termed by Granger and Newbold (1974).[5] The increasingly important role played by panel data in the time series analysis makes the panel unit root test a standard procedure. If the variables in equations (4) and (5) are not stationary, a panel cointegration technique may be needed to examine the cointegration relationship among the variables.

A primary purpose for using the panel unit root is to enhance the power of the test due to its large sample property. Even though such a property can also be obtained by increasing the length of the sample period, a long time period may well suffer from structural changes. On the other hand, the cross section property of panel data introduces heterogeneity and hence weakens the testing power. In the initial stage of the panel unit root models, heterogeneity is not allowed for (e.g., Levin and Lin, 1993; Breitung and Meyer, 1994; and Quah, 1994). A recent advance by Im, Pesaran, and Shin (IPS, 1997) takes cross-section heterogeneity into consideration. We briefly describe the IPS model:

Given y_{it} (i = 1, …, N and t = 1, …, T), we generate $\tilde{y} = y_{it} - (\frac{1}{N}) \sum_{i=1}^{N} y_{it}$, and apply the

following model similar to the Augmented Dickey-Fuller test (ADF) on $\tilde{y}_{it}$

$$\Delta \tilde{y}_{it} = \alpha_i + \beta_i \tilde{y}_{it-1} + \sum_{j=1}^{w_j} \gamma_{ij} \Delta \tilde{y}_{it-j} + \varepsilon_{it} \qquad (6)$$

We use t_{ij} (w_i) to test the null of (6) $\hat{\beta}_i = 0$ and define the t – bar statistic from the following relation:

$$\bar{\Gamma}_{\bar{t}} = \sqrt{N}\,(\bar{t}_{NT} - a_{NT}\,)/\sqrt{b_{NT}} \quad \sim \quad N(0,1)$$

where $\quad \bar{t}_{NT} = \dfrac{1}{N}\sum_{i=1}^{N} t_{iT}\,(w_i\,),$

$$a_{NT} = (\frac{1}{N})\sum_{i=1}^{N} E[\,t_{iT}\,(w_i)\,/\,\beta_i = 0\,],\text{ and} \tag{7}$$

$$b_{NT} = \dfrac{1}{N}\sum_{i=1}^{N} V[\,t_{iT}\,(w_i)\,/\,\beta_i = 0\,].$$

Note that a_{NT} and b_{NT} in (7) need to be estimated before applying the IPS test. Table 2 of their paper provides the simulated values from the Monte Carlo experiments. The testing power is shown to be superior to that by Levin and Lin (1993). The t statistics from (6) for all states are then substituted into (7) to obtain the t–bar statistics. A comparison with a 5-variable critical value at $\alpha = 5\%$ is shown in Table 1.

A perusal of Table 1 indicates the panel variables are stationary. As such, we employ the pooling technique (fixed effect model) to estimate the demand structure of (4).

Discussion of the Empirical Results

The advantage of pooled estimators over heterogeneous estimators (individual state demand functions) is that individual regressions often yield unreliable and implausible coefficients. For example, long run price elasticity using OLS varies from 5.46 to –8.46 and from 5.23 to –7.46 using 2SLS (Baltagi *et al*., 2000) despite that Pesaran and Smith (1995) showed the average of these estimates (mean group or MG estimator) is consistent as N and T approach infinity. Besides, the MG estimator is sensitive to outliers (Pesaran *et al*., 1999, p. 629). Other heterogeneous estimators such as shrinkage methods by Maddala *et al*. (1994) produce equally unstable estimates (Baltagi *et al.,* 2000, p. 122). This reduces the choice to the two remaining estimators: the dynamic fixed-effect (DFE) models in which parameter homogeneity prevails except intercepts and the pooled mean group (PMG) model where intercepts, short-run elasticities, and error variance can vary but the long-run price elasticity is constrained to be the same for all states. While it seems true that income elasticity of consumption should approach 1 in the long run for a country according to Pesaran *et al*. (2000), the price elasticity may not converge to a common value even in the long run for an addictive commodity. Such a non-convergent and volatile demand structure could be traced to the switching elasticity theorem by Greenhut *et al*. (1974). As the phenomenon of declining price elasticity at higher prices has been borne out in many empirical studies (e.g., Keeler *et al*., 2001), its role cannot be downplayed.

In the panel data estimation, one may choose between the random effect and fixed effect models. According to Judge *et al*. (1988), the difference in results is miniscule for a large T and a small N. In the case of a small T relative to N, the fixed-effect model is inefficient though consistent and, as such, a random effect model may be preferred. However, the correlation between characteristics pertaining to cross sections and explanatory variables may render estimates biased in the random effect model. For a relatively smaller T (as T = 42, N = 43 in our case) the worst in applying the fixed effect model is its inefficiency. Above all, 41 regions

constitute nearly 85% of the entire U.S. and as such, treating parameters as fixed is a reasonable assumption.

Besides, Baltagi *et al*. (2000) and Freeman (2000) pointed out that the fixed effect model performs better in prediction among the pooling techniques. Consequently, we apply the fixed effect model to estimate (4). The choice of the fixed effect model is also confirmed by the Hausman specification test.[6] Since the cigarette price is determined by both demand and supply, an OLS estimate of (4) could very likely lead to a simultaneous equation bias. Notice that the lagged consumption (Q_{it-1}) is considered an endogenous variable for it is related to the error term. As such, we opt for the two-stage least square (2SLS) model with the instruments of real cigarette tax, elderly population proportion ages 65 years and older, lagged values of real cigarette tax, cigarette price, neighborhood price, per capita sale, and per capital income.[7] Results are shown in Table 2.

An inspection of Table 2 indicates a price, income, and substitution elasticity of -0.41, 0.06, and 0.3 respectively with expected signs and significant t values. The coefficient on the elderly population proportion appears to be statistically significant with the correct sign: as the elderly population proportion increases, cigarette consumption is expected to decrease. Note that the OLS estimates differ from that of the 2SLS which takes the endogeneity problem into consideration. Our 2SLS estimates (1954-1997) are very similar to that by Keeler *et al.* (2001) in which the omitted variable is considered for the sample period of 1960-1990 and panel estimate results by Yurekli and Zhang (2000). In fact, prior panel results are rather diverse. For example, the price and income elasticities by Baltagi *et al.* (2000), who used 1963-1992 panel data are in the range of -0.009 ~ -0.48 and -0.04 ~ 0.19. The elasticities are found to be between -0.47 ~ -0.61 and -0.0001 ~ 0.0002 by Yurekli and Zhang, who used 1970-1995 panel data. Clearly, the results vary with respect to sample lengths. As more evidence has surfaced against smoking, it is of interest to investigate the size of these elasticities. To this end, we make use of the rolling window approach in estimating the elasticities. That is, for a window length of 15 years, we first estimate the 1975 elasticities via the window period of 1961-1975. For the 1976 elasticities, we use the window of 1962-1976. Finally, the 2002 elasticities are estimated via the 1988-2002 windows. Both price and income elasticies, short run and long run, for the window length of 20 and 25 years are estimated and shown in Figure 1.

A perusal of Figure 1 suggests that short run price elasticity values are between -0.25 and -0.7 for all three windows. Roughly speaking, the longer the length of the window is, the less volatile the estimated price elasticity will be. Furthermore, it appears that the price elasticity is on the decline (absolute value). Moreover, volatility of the income elasticity is also in decline, most likely due to increasing alertness about harmful cigarette smoking. In sum, we have witnessed declining values for both elasticities as more smokers quit smoking; the remaining "hardcore" consumers represent the "trapped" price inelastic group. Beyond that, cigarette consumption tends to become an inferior good. Note that estimated time-varying elasticities seem to become stable once the sample length reaches 25 years. For instance, the rolling elasticity values are sensitive to the end period (1992 or 1995) if 15-year or 20-year sample data are used. Prior estimates on price elasticity, -0.2581 by Thursby and Thursby (2000) and -0.49 ~ -0.61 by Yurekli and Zhang (2000), vary noticeably once the sample length is not long enough. Based on our simulations, it seems a reliable estimate on elasticities entail the use of a longer sample length (*e.g.,* at least 25 years) in estimating cigarette demand.

Long-run price and income elasticities (-2.5 ~ -1.0) and (0.1 ~ 0.8) are similar as they largely mirror the short-run elasticities via the coefficient of lagged consumption (ρ). To dissect

the price elasticity of cigarettes, we need to divide its price into pre-tax price PT_{it} and cigarette tax Tax_{it} deflated by CPI. The resulting demand can then be estimated based on the following formulation:

$$\ln Q_{it} = \alpha_1 + \rho \ln Q_{it\text{-}1} + \beta_{11} \ln PT_{it} + \beta_{12} \ln Tax_{it} + \beta_2 \ln Y_{it} + \beta_3 \ln Pn_{it}$$
$$+ \beta_4 T + \beta_5 \, Age \, 65_{it} + \mu_{it} \tag{8}$$

The time-varying pre-tax and tax elasticities of 20-year and 25-year moving windows estimated by the instrument variable approach are presented in Figure 2.

As with price elasticity of demand in which expenditure-income share is an important determinant, real tax / real price ratio must be considered in explaining demand price elasticity of cigarettes. Let us denote gross price and tax by P and t. Hence, the net price PT is

$$PT = P - t . \tag{9}$$

A few algebraic manipulations lead readily to

$$\varepsilon_P = \frac{\varepsilon_{PT} \cdot \varepsilon_t}{(1-\alpha)\varepsilon_t + \alpha\varepsilon_{PT}} , \tag{10}$$

where ε_P and ε_{PT} denote price and pre-tax elasticity, ε_t is tax elasticity; and α is tax share shown as a percentage to price or t/P. In general, tax elasticity decreased in absolute value when tax share α decreased until about 1992 (Figure 2 (a)). In contrast, cigarette consumption was relatively more tax elastic when the tax share was higher in most of the 1970's. With a greater tax share, cigarette consumption is more sensitive to the tax price. In terms of the denominator of equation (10), for each increase in α, the decrease in $(1-\alpha)\varepsilon_t$ is greater than an increase in $\alpha\varepsilon_{PT}$: tax elasticity is generally greater than pre-tax price elasticity in absolute value (Figures 2a and 2b), i.e., $|\varepsilon_t| > |\varepsilon_{PT}|$ for a stable numerator. The smaller denominator in (10) signals a relatively price elastic cigarette demand around 1974 ~ 1975 (Figure 1). On the other hand, a declining tax share (α) gave rise to a greater denominator in (10). Hence, the price elasticity became much more inelastic around 1992 when the tax share bottomed out to about 25% of the price. That is, the cigarette taxes were relatively small during the period.

Note that equation (10) is derived in deterministic form for small change whereas the time-varying rolling-window elasticties are statistical estimates with random variations. Equation (10) provides simulated elasticity that generally fits estimated elasticity, especially when real price changes are moderate, i.e., -0.025 (simulated value based on (10)) versus –0.029 (rolling window value) for 1995; -0.04 versus 0.04 for 1996; and –0.036 versus 0.044 for 1997.

Income elasticity peaked to 0.24 around 1980 and then gradually plummeted to 0.02 in 1995 (Figure 1). The declining trend can be better understood by disaggregating income (Y_{it}) into earning $(EARN_{it})$, dividend (Dvd_{it}), and transfer payment (Trp_{it}) components.[8] The three income elasticities within the 20-year (Figure 3a) and 25-year (Figure 3c) rolling window frameworks along with the corresponding income shares (Figures 3b and 3d) are shown in Figure 3.[9] An examination of Figure 3a and Figure 3c indicates that both earning elasticity and transfer

income elasticity were on the rise before 1992. On the other hand, the dividend income elasticity after 1992 was in sharp decline (25-year window) as was the transfer income elasticity after 1990. Generally speaking, decreasing income elasticity after 1992 (Figure 1b) can be attributed to its three components. Dividend income elasticity had the largest decline, followed by transfer income elasticity and earning elasticity (Figure 3). After 2000, the earning elasticity was actually on a mild rise, due to a recession as cigarette consumption may well be considered counter-cyclical. In short, decreases in both dividend income and transfer income elasticities account to a large degree for the declining income elasticity in the U.S.

As indicated by Keeler *et al*. (1993), cigarette consumption is related to the education level. Generally speaking, a better educated group is more accessible and acceptable to the information of severe health problems from smoking. Beyond that, there appears to be a positive relationship between income and education level. Dividend, as well as interest income, can be viewed as a proxy for the income of that group with more education. In this light, it becomes evident that cigarette smoking is an inferior good for the higher income group. Furthermore, the information about smoking from the Surgeon General in 1986, 1988, 1990, and 1992 has also discouraged smoking to some extent, even for wage income and transfer payment groups. The declining income elasticity can be attributed to (i) increasing income share for dividend recipients and decreasing share for wage earners, (ii) decreasing income elasticities from a sequence of warnings about hazardous health especially in the 1990's, (iii) negative income elasticity for a dividend recipient group, and (iv) declining transfer income elasticity with the principal recipients of the elderly and the poor, who cannot afford more cigarette consumption in the presence of rising prices. This negative relation between the elderly population and cigarette consumption can also be detected in Table 2.

Concluding Remarks

Undoubtedly, the Surgeon General's report on harmful effects caused by cigarette smoking has affected its demand structure. A tax policy could be effective for forward-looking consumers in reducing consumption. Based on the Markov Switching model, Coppejans and Sieg (2002) estimated about 31% of smokers in Los Angeles believe the future price will rise by at least 10%. The relative tax – which went up close to 50% in the beginning but plummeted since the 1970's – is on the rise again. As litigation against big tobacco companies continues to favor the plaintiff, the ability to raise tobacco price plays a critical role. That is, only a very inelastic demand in the future can essentially provide a reliable source for the astronomical amount of settlement costs. This paper makes use of the panel unit root technique recently developed by Im *et al*. (1997) to test the spurious regression problem. Considerable amount of data – 42 states and Washington, D.C. spanning from 1961 to 2002 – are assembled via the fixed effect model to yield the price, income, and substitution elasticities of cigarettes of –0.41, 0.06, and 0.30, respectively. Three interesting results emerge from the panel estimates of the cigarette demand structure. First, this paper using 15-year, 20-year, or 25-year rolling window approaches finds decreasing price (absolute value) and income elasticities. In addition, we show that demand price elasticity is dependent largely on tax share and tax elasticity. The decreasing tax shares contribute to smaller tax elasticity (absolute value), which in turn gives rise to a more inelastic demand. Given the rising state budget deficits, cigarette taxes will increase as will the cigarette price owing to the enormous size of settlement costs imposed on the tobacco industry.

Last, the declining income elasticity can be better analyzed by breaking income down into three categories. The time varying elasticities of individual groups indicates (i) anti-smoking campaigns have contributed to declining income elasticity; (ii) increasing income shares for dividend recipients and decreasing income share for wage earners generally account for decreasing income elasticity because the former are more likely to quit or reduce cigarette smoking; (iii) for dividend recipients, cigarette consumption has recently become an inferior good; and (iv) decreasing transfer income elasticity. Therefore, future cigarette consumption will critically hinge on tax share, income share, price, and income elasticities of remaining hard-core smokers.

Footnotes

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1. Variables such as beverage, alcohol and bubble gum consumption could have been included in Z_{it}, if available.

2. The partial adjustment model still remains the main framework in studies by Keeler *et al.* (2001), Yurekli and Zhang (2000), and Thursby and Thursby (2000) since the estimated discount rates in the rational addiction model are not feasible or negative.

3. Forty-two states are included in the study owing to the fact that neighboring state prices, and/or elderly population proportion data are not available. The missing eight states are Alaska (AK), Colorado (CO), Hawaii (HI), Maine (ME), North Carolina (NC), New Hamshire (NH), Oregon (OR), and Washington (WA).

4. ADF statistics of average income, average cigarette price, and consumption for 42 states and Washington, D.C. are −1.42, -0.89, and −0.92, respectively with the corresponding critical value of −3.52, -3.52, and −2.93, respectively: an indication of the existence of an unit root without using panel data.

5. The spurious regression leads to estimation bias due to non-standard distribution, hence the correlation estimation may well be biased.

6. We employ the Hausman (1978) specification test with the χ^2 statistic (d. f. = 5) of 39.96 to reject the null hypothesis. Hence, we do not use the random effect model.

7. One lag period is used.

8. Official definitions of the three income groups are due to Current Population Survey and the U.S. Census Bureau.

9. Figures 3b and 3d contain the same information since income shares are not estimated under either window length.

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Table 1. Results of the Panel Unit Root Test

Variables	$\overline{\Gamma}_\iota$
Q_{it}	-2.82*
P_{it}	-2.11*
Pn_{it}	-2.31*
Y_{it}	-4.25*

Note: The 5% critical value, for a_{NT} = 1.470, b_{NT} = 0.801 (with one lag and time trend) is −1.96. The null hypothesis is: the variable has a unit root. ∗ = 1% significance level. ** = 5% significance level.

Table 2. Panel Estimates of the Cigarette Demand

	$\ln Q_{it}$	$\ln P_{it}$	$\ln Y_{it}$	$\ln Pn_{it}$	T	$Age65_{it}$	$\overline{R}^2$
OLS	0.8874 (48.76)	-0.2114* (-7.75)	0.0421 (1.38)	0.1322* (5.37)	-0.0016** (-2.26)	-0.0032*** (-1.67)	0.98
TSLS	0.8490 (42.31)	-0.4067* (-8.81)	0.0579*** (1.64)	0.3036* (7.62)	-0.0019** (-2.31)	-0.0046** (-2.34)	0.98

Note: Numbers in parentheses are t statistics; * = 1% significance level. The estimation is based on the fixed-effect model. Standard errors are hetero-scedasticity-consistent estimator by White. Q_{it} = tax-paid per capita sales for state i and time period t (subscripts); P_{it} = price per pack deflated by CPI; Y_{it} = personal income deflated by CPI; Pn_{it} = average price of adjacent states deflated by CPI; and T = time trend (1961= 1). Age 65 = proportion of population ages 65 years and older in the ith state.

Figure 1. Time Varying Elasticities of Cigarette Demand

(a) Time Varying Short Run Price Elasticities

—— 15 Years Rolling Window
----- 20 Years Rolling Window
——— 25 Years Rolling Window

(b) Time Varying Short Run Income Elasticities

—— 15 Years Rolling Window
----- 20 Years Rolling Window
——— 25 Years Rolling Window

(c) Time Varying Long Run Price Elasticities

—— 15 Years Rolling Window
----- 20 Years Rolling Window
——— 25 Years Rolling Window

(d) Time Varying Long Run Income Elasticities

—— 15 Years Rolling Window
----- 20 Years Rolling Window
——— 25 Years Rolling Window

Figure 2. Tax and Pre-Tax Elasticities and Tax Share(20-year rolling window)

(a) Time Varying Pre-Tax Price and Tax Elasticities

——— Real Tax as % of Real Price
------- 20 Years Rolling Window - Pre Tax Price
———— 20 Years Rolling Window - Tax
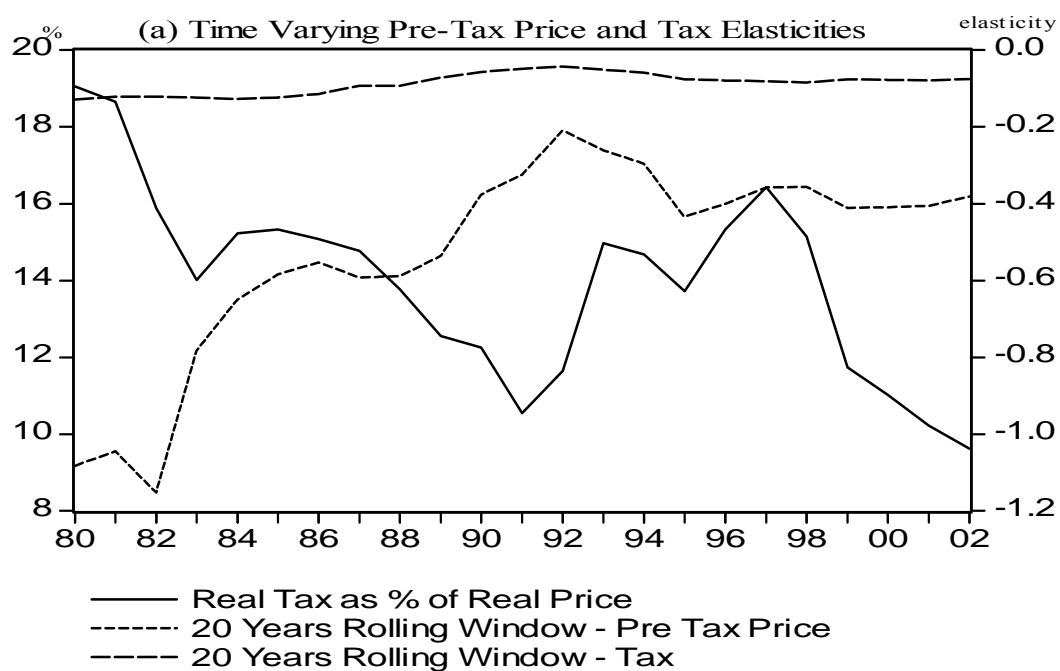
Tax and Pre-Tax Elasticities and Tax Share(25-year rolling window)

(b) Time Varying Pre-Tax Price and Tax Elasticities

——— Real Tax as % of Real Price
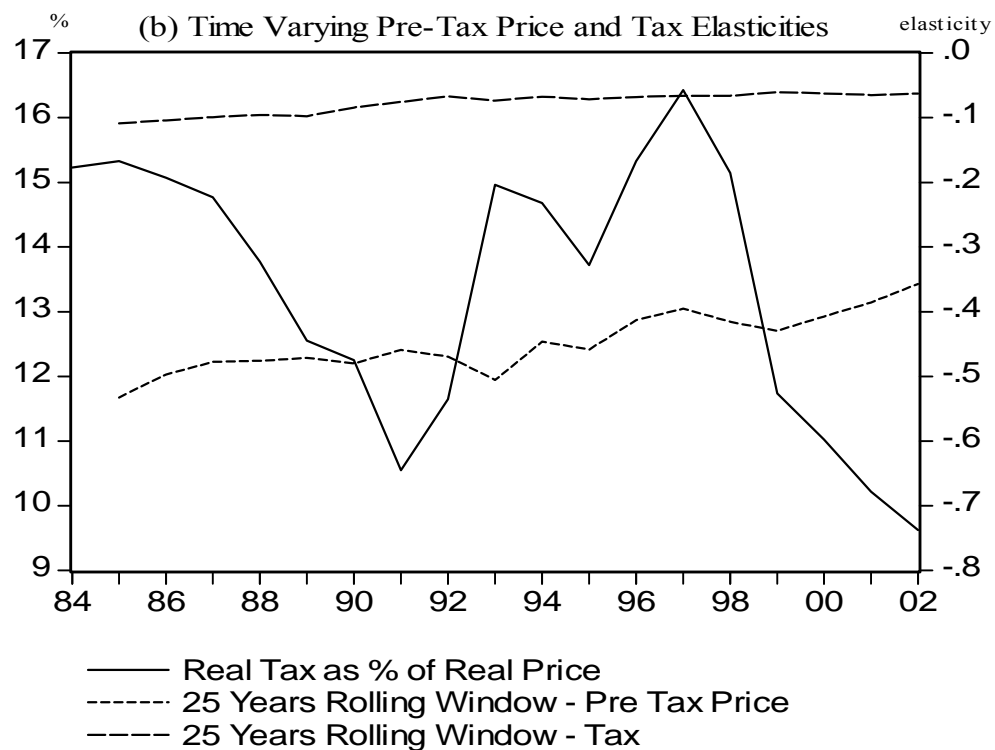------- 25 Years Rolling Window - Pre Tax Price
–––– 25 Years Rolling Window - Tax

Figure 3. Component Income Elasticities and Income Shares

20-year rolling window

(a) Time Varying Income Component's Elasticities

──── 20 Years Rolling Window - Earning
----- 20 Years Rolling Window - Dividend
── ── 20 Years Rolling Window - Transfer Payment

(b) Time Series of Relative Income Shares

──── Earning as % of Income
----- Dividend as % of Income
── ── Transfer Payment as % of Income

25-year rolling window

(c) Time Varying Income Component's Elasticities

──── 25 Years Rolling Window - Earning
----- 25 Years Rolling Window - Dividend
── ── 25 Years Rolling Window - Transfer Payment

(d) Time Series of Relative Income Shares

──── Earning as % of Income
----- Dividend as % of Income
── ── Transfer Payment as % of Income