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Advertising and Cigarette Consumption

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INTRODUCTION

Whether a ban on all types of cigarette advertising would reduce cigarette demand is a central issue in the recent debate over imposing such a policy. Health organizations seem to suppose that such a ban would lower demand: the recent proposal by the American Medical Association that all cigarette advertising be banned was quickly endorsed by the American Heart Association, the American Lung Association, and the American Cancer Society. However, the majority opinion of economists is summed up in the *Economic Report of the President* [28]: "there is little evidence that advertising results in additional smoking." This hypothesis is supported by numerous studies including Baltagi and Levin [3], Hamilton [9], Schmalensee [22], and Schneider *et al.* [23].

The minority view that advertising increases consumption is supported by Bishop and Yoo [4], Fujii [8], and Kao and Tremblay [15] for the U.S.; McGuinness and Cowling [18], Radfar [20], and Witt and Pass [25] for the U.K.; Leu [16] for Switzerland, and McLeod [19] for Australia. But these studies have shortcomings. McGuinness and Cowling [18], Radfar [20], and Witt and Pass [25] use incomplete estimates of advertising expenditures. Leu [16] and McLeod [19] have no advertising data at all; they use advertising ban variables as proxies for advertising. Fujii [8] uses wholesale price rather than retail price plus tax. The results of Bishop and Yoo [4] are questionable on three counts. First, they implicitly assume that advertising effects depreciate within one year. If this assumption is incorrect then their advertising coefficient could be biased. Second, Bishop and Yoo [4] estimate a supply and demand system. But a supply function does not exist in an industry with market power. In fact, the results of Ashenfelter and Sullivan [1] support Schmalensee's [22] contention that price competition is absent in the cigarette market.[1] Third, Bishop and Yoo [4] do not control for population despite their use of aggregate consumption instead of the more common *per capita* measure. Kao and Tremblay [15] correct the last two points but maintain the assumption that advertising depreciates within a year. While the assumption is plausible, we view it as a hypothesis to be tested.

Baltagi and Levin [3] study consumption in the various states over the period 1963–80, alternatively assuming that advertising has an infinite lag and that the lag is finite. Our study differs from theirs since we use aggregate data over a longer time period.

We examine two policy questions: (1) is there a positive response of aggregate demand for cigarettes to advertising? (2) what is the reaction of consumers to government health warnings and media policy? Our data cover the period 1952–84, so our time series is longer than any previously used with the exception of one of the series in Hamilton [9], and includes more recent data than other studies. Since we use such current data, we include a binary variable to estimate the effect of the 1979 Surgeon General's report which expanded upon the 1964 Surgeon General's report.

Our statistical results support the minority hypothesis that cigarette advertising increases aggregate demand. The 1964 and 1979 government health warnings and the 1968 and 1971 media policies are

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all found to reduce the aggregate demand for cigarettes. The findings also support the hypothesis that advertising effects depreciate within one year.

THE DEMAND EQUATION

The form of the demand equation used in the regression equation is similar to the habit persistence models of Baltagi and Levin [3], McGuinness and Cowling [18], and Radfar [20]. Unlike these studies, we endogenously estimate the depreciation rate of advertising. We follow these studies by treating the health warning and media policy dummy variables as slope dummies associated with the advertising coefficient rather than as intercept dummies as in earlier studies. The slope dummies suggest that health warnings and media policies affect consumption through the advertising elasticity. The dummy variables are for the 1964 Surgeon General's report (D64 = 1 for years 1964–present, 0 otherwise), for the effective years of the Fairness Doctrine[2] (D68 = 1 for years 1968 through 1970, 0 otherwise), for the broadcast advertising ban (D71 = 1 for years 1971–present, 0 otherwise), and for the 1979 Surgeon General's report (D79 = 1 for years 1979–present, 0 otherwise). We assume that the that the "desired" *per capita* consumption level Q_t^* is given by

$$\ln Q_t^* = \alpha_0 + \alpha_1 \ln P_t + \alpha_2 \ln Y_t + [\alpha_3 + \mu_1(D64_t) + \mu_2(D68_t) + \mu_3(D71_t) + \mu_4(D79_t)] \ln G_t$$

where Q_t^* is in thousands of cigarettes, P_t is the real retail price per thousand (including taxes), Y_t is real *per capita* disposable income, $D64_t$, $D68_t$, $D71_t$, and $D79_t$ are the dummy variables, and G_t is the *per capita* demand shifter (which we call "goodwill") created by current and past advertising.[3]

Consumption does not adjust immediately to the desired level. We model changes in consumption as a partial adjustment process:

$$\ln Q_t - \ln Q_{t-1} = \pi(\ln Q_t^* - \ln Q_{t-1})$$

where Q_t is the actual *per capita* consumption (in thousands of cigarettes)[4] per year and π is the adjustment (or habit persistence) coefficient. Substituting $\ln Q_t^*$ into the adjustment equation, rearranging terms, and adding a normal error term, u_t, to the result yields

$$\text{(1)} \quad \ln Q_t = \pi\alpha_0 + \pi\alpha_1 \ln P_t + \pi\alpha_2 \ln Y_t + \pi[\alpha_3 + \mu_1(D64_t) + \mu_2(D68_t) + \mu_3(D71_t) + \mu_4(D79_t)] \ln G_t + (1-\pi)\ln Q_{t-1} + u_t.$$

The coefficients are elasticities. For example, $\pi\alpha_1$ is the short-run price elasticity of demand while α_1 is the long-run (full adjustment) price elasticity of demand.

We assume that *per capita* goodwill changes as follows:

$$\text{(2)} \quad \Delta G_t = G_t - G_{t-1} = A_t - \delta G_{t-1}$$

where Δ is the difference operator, A_t is real *per capita* advertising expenditures at time t, and $\delta \in [0,1]$ is the depreciation rate for goodwill. Rearranging terms in equation (2) and substituting recursively for G_{t-j}; j = 1, 2, . . . ; yields

$$\text{(3)} \quad G_t = A_t + (1-\delta)A_{t-1} + (1-\delta)^2 A_{t-2} + \cdots + (1-\delta)^m A_{t-m} + \cdots$$

If the advertising terms beyond lag m are small enough to ignore, we can substitute equation (3) into equation (1) to obtain

$$\text{(4)} \quad \ln Q_t = \pi\alpha_0 + \pi\alpha_1 \ln P_t + \pi\alpha_2 \ln Y_t + \pi[\alpha_3 + \mu_1(D64_t) + \mu_2(D68_t) + \mu_3(D71_t) + \mu_4(D79_t)] \times \ln\{A_t + (1-\delta)A_{t-1} + (1-\delta)^2 A_{t-2} + \cdots + (1-\delta)^m A_{t-m}\} + (1-\pi)\ln Q_{t-1} + u_t.$$

STATISTICAL RESULTS

Equation (4) is estimated using nonlinear ordinary least squares (NOLS) over the period 1952–84 using annual data. Data are discussed in Appendix A. We also tried a remainder term as in Baltagi and Levin [3], but the coefficient on the remainder term always took the wrong sign. For this reason, and because we find advertising to depreciate very quickly, we do not report those results. Since equation (4) treats price as an exogenous variable, we report results of a Hausman test which supports the exogeneity of price. Price exogeneity is plausible because price competition appears to be absent in the cigarette market.

Given the functional form of equation (4), the number of advertising lags (m) is necessarily limited. The largest value of m which we use is 10. Previous studies which consider lagged advertising have achieved their best results with less than 10 lags (*e.g.,* Baltagi and Levin [3] and McGuinness and Cowling [18]).

Since the demand equation contains a lagged dependent variable, the presence of autocorrelation is tested using Durbin's [6] $\hat{K}_2 \sim \chi^2(1)$.[5] The null hypothesis is that the autocorrelation coefficient is zero. In our case, the critical value at the 5 percent level is 3.84. If $\hat{K}_2 > 3.84$ then we correct for autocorrelation by estimating the model's coefficients and the autocorrelation coefficient simultaneously. Details are given in Appendix B.

Results are presented in Table 1 for regressions which include advertising lags of 0, 1, 2, 4, 6, 8, and 10 periods.[6] The zero-lag case assumes m = 0 or, equivalently, $\delta = 1.0$. Autocorrelation seemed to be a problem in the equations with 6, 8, and 10 lags, so the correction procedure was applied in these regressions.[7] The estimates and *t* values of the autocorrelation coefficient ρ in Table 1 for these equations indicate negative autocorrelation.

The results in Table 1 lead us to focus attention on the model with no lagged advertising terms for two reasons. First, the depreciation term (δ) exceeds unity in the equations with lagged advertising terms and tends to increase with the number of lagged terms beyond four. It is meaningless for the depreciation term to exceed one, yet this is consistent with Baltagi and Levin [3] who find that their regressions improve as the rate approaches one. Second, the estimates are all significant in the equation with no lagged advertising and the elasticities are reasonable.[8]

We use Hausman's [10] *m* statistic to test price endogeneity in the equation with no lagged advertising terms. We calculate $m = \hat{q}'\hat{B}^{-1}q'$. q' is a column vector with elements equal to the difference between coefficient estimates from our NOLS regression and coefficients from a nonlinear instrumental variable (NIV) regression. The NIV equation is identical to the NOLS equation except that the price variable is replaced by an instrument constructed by regressing the price variable on a set of exogenous variables.[9] (Details regarding the instrument and the Hausman test are available from the authors upon request.) $\hat{B}$ is the matrix of differences between the covariance matrices of the NOLS and NIV equations multiplied by the inverse of the number of data points. The null hypothesis is that the NOLS model is correctly specified. In our case, $m \sim \chi^2(9)$ with critical value 14.6837 at the 10 percent level. We calculate $m = 3.4979$. Therefore, we cannot reject the hypothesis that the NOLS model is correctly specified, and we need not endogenize price.

In the equation with no lagged advertising term, the short-run price elasticity ($\pi\alpha_1$) is -0.10 with a *t*-ratio of -1.352. This is significant at the 10 percent level in a one-tailed test.[10] Our estimate is smaller in absolute value than the estimate of -0.20 obtained from Baltagi and Levin [3], but this is expected. They use state-level data, which allows individuals to substitute bootleg cigarettes from another state. Our demand function is more highly aggregated so there are no substitutes. Estimates in Bishop and Yoo [4], Kao and Tremblay [15], Schmalensee [22], and Schneider *et al.* [23] range from -0.406 to -1.218. While our price elasticity is smaller in absolute value than these estimates, it supports the popular notion of a highly price inelastic demand curve for cigarettes due to the habitual nature of cigarette consumption.

Our long-run price elasticity (α_1) is -0.32 with *t*-ratio -1.38, significant at the 10 percent level in a one-tailed test. Baltagi and Levin [3] suggest a long-run elasticity of -0.22 while Schmalensee [22] suggests a value of -0.46. The differences among these values are due in part to different habit persistence coefficients. Our lagged consumption coefficient ($1-\pi$) is nearly 0.70 and is close to the estimate of Schmalensee [22], which is 0.71.

Our short-run income elasticity is 0.29 with *t*-ratio 3.032 and is close to Schmalensee's [22] estimate

TABLE 1
Non-Linear OLS Estimates of Coefficients of Cigarette Demand, 1952–1984

Coefficients	Zero Lag	One Lag	Two Lags	Four Lags	Six Lags	Eight Lags	Ten Lags
α_0	−5.557	−3.729	−4.418	−0.631	2.421	2.292	2.225
	(−2.15)**	(−1.92)*	(−1.98)*	(−0.21)	(1.71)*	(3.21)***	(2.78)***
π	0.318	0.368	0.340	0.260	0.458	0.796	0.726
	(2.94)***	(3.58)***	(3.00)***	(2.51)**	(4.29)***	(4.56)***	(3.69)***
α_1	−0.323	−0.410	−0.331	−0.476	−0.383	−0.319	−0.391
	(−1.38)	(−2.03)**	(−1.48)	(−1.78)*	(−4.23)***	(−7.28)***	(−4.22)***
α_2	0.898	0.730	0.773	0.417	0.038	0.023	0.063
	(2.87)***	(3.14)***	(2.87)***	(1.26)	(0.26)	(0.27)	(0.59)
α_3	0.201	0.261	0.226	0.086	−0.022	0.001	−0.022
	(2.63)***	(3.56)***	(3.02)***	(0.76)	(−0.43)	(0.01)	(−0.40)
μ_1	−0.128	−0.166	−0.139	−0.092	−0.007	−0.005	−0.004
	(−2.46)**	(−2.94)***	(−2.54)***	(−1.53)	(−0.31)	(−0.40)	(−0.28)
μ_2	−0.073	−0.091	−0.078	−0.060	−0.038	−0.030	−0.029
	(−2.10)**	(−2.23)**	(−2.08)**	(−1.41)	(−2.09)**	(−3.80)***	(−2.64)***
μ_3	−0.146	−0.223	−0.159	−0.119	−0.070	−0.059	−0.064
	(−2.34)**	(−3.68)***	(−2.34)**	(−1.60)	(−2.16)**	(−3.95)***	(−3.35)***
μ_4	−0.069	−0.070	−0.076	−0.068	−0.049	−0.055	−0.054
	(−2.11)**	(−1.84)*	(−2.02)**	(−1.51)	(−2.34)**	(−9.78)***	(−9.07)***
δ	(See +	1.482	1.524	1.388	1.445	1.875	1.943
	below)	(12.54)***	(2.12)**	(1.86)*	(2.60)***	(11.34)***	(23.06)***
ρ	—	—	—	—	−0.502	−0.700	−0.743
					(−1.84)*	(−2.91)***	(−3.43)***

The following levels of significance are for a two-tailed test:
* significant at 10% level.
**significant at 5% level.
***significant at 1% level.
+When the lag is set at zero, lagged advertising terms drop out, so δ implicitly equals 1.

Note: *t*-ratios are shown in parentheses.

Note: Equations with 6, 8, and 10 lags have been corrected for serial correlation. ρ is the estimated serial correlation coefficient.

of 0.32. Previous estimates range widely from −0.004 (Baltagi and Levin [3]) to 1.482 (Kao and Tremblay [15]). Our result supports the intuitively plausible notion that a change in income will have little impact on consumption.

$G_t = A_t$ when no lagged advertising terms is included. Our short-run advertising elasticity ($\pi\alpha_3$) is 0.06 with a *t*-ratio of 2.293 while the long-run advertising elasticity (α_3) is 0.20 with a *t*-ratio of 2.63. This suggests that advertising could increase demand in the absence of government health warnings and media policy. The coefficient for the 1971 advertising ban is significant and negative. Coupled with the significant advertising elasticity, this implies that a more complete advertising ban would further reduce cigarette consumption. However, further investigation is warranted. We will analyze the combined short-run effects of advertising and government policy; the long-run analysis is similar.

For the periods 1964–67, 1968–70, 1971–78, and 1979–84 the complete advertising coefficients are $\pi(\alpha_3 + \mu_1)$, $\pi(\alpha_3 + \mu_1 + \mu_2)$, $\pi(\alpha_3 + \mu_1 + \mu_3)$, and $\pi(\alpha_3 + \mu_1 + \mu_3 + \mu_4)$.[11] The estimates (*t* values) are

0.0232 (0.8755), 0.0001 (0.0140), −0.0233 (−0.615), and −0.0452 (−1.337). The first three estimates are small and insignificant at the 10 percent level, indicating that government health warnings and media policies have rendered advertising incapable of effectively increasing demand. The last elasticity is negative and significant in a one-tailed test at the 10 percent level.

A negative estimate for the advertising coefficient is quite plausible. One can think of a health warning or media policy as generating anti-smoking "signals". Let S_t be the level of these signals and suppose that $\sigma S_t = A_t$ where $\sigma > 0$.[12] Omitting lagged advertising terms, the demand equation with anti-smoking signals is

$$(5)\quad \ln Q_t = \pi\alpha_0 + \pi\alpha_1 \ln P_t + \pi\alpha_2 \ln Y_t + \pi[\mu_1 (D64_t) + \mu_2(D68_t) + \mu_3(D71_t) + \mu_4(D79_t)] \ln \sigma S_t + \pi\alpha_3 \ln A_t + (1 - \pi) \ln Q_{t-1} + u_t.$$

But since $\sigma S_t = A_t$, equation (5) is precisely equation (4) which is the equation that we estimate. From equation (5) and the statistical results, industry advertising appears to increase demand while the same advertising also generates signals which decrease demand.[13]

CONCLUSION

Our statistical results indicate that advertising increased the aggregate demand for cigarettes during the period 1952–63. After 1963, government health warnings and media policies seem to have eliminated the aggregate demand shifting potential of the advertising.[14] This supports the intuition of many health organizations that advertising increases cigarette consumption while advertising bans decrease cigarette consumption. We may not, however, conclude that advertising bans involving all media would further reduce cigarette consumption. Our results suggest that past government policies have already effectively reduced the aggregate advertising elasticity to zero, so that at present advertising is simply redistributing market share. Therefore, the additional benefits of a total advertising ban may be small while the costs may or may not be small.[15]

Most previous cigarette demand studies fail to obtain a statistically significant coefficient for advertising. One might wonder why our results should differ. The difference may be due to the length of our data set and our choice of regression method. Data associated with cigarettes are notoriously collinear.[16] Perhaps collinearity is responsible for the apparent statistical insignificance of advertising in previous studies. Two ways to mitigate collinearity problems are to increase the data set and to impose reasonable restrictions on the coefficients. Our data set is longer than others[17] and NOLS allows us to impose cross-coefficient restrictions upon the relationships among the coefficients.

APPENDIX A

Definitions and the Sources of Data

Q = thousands of cigarettes *per capita*. Total quantity is from Maxwell [17] and smoking age population is U.S. population 14 years and over from *Historical Statistics of the United States, Colonial Times to 1970* and various issues of the *Statistical Abstract of the United States*.

P = real retail price per thousand cigarettes. Calculated from the median nominal retail price per pack (including taxes) for all taxing states which is reported in *The Tax Burden on Cigarettes* [24]. Nominal price is divided by the 1980-based consumer price index in *International Financial Statistics* [12].

Y = real *per capita* disposable income. Reported in the *Economic Report of the President* [28] and adjusted by the 1980 consumer price index.

A = real *per capita* advertising expenditures. The series for total advertising expenditures was constructed from three sources. The Federal Trade Commission [7] reports advertising data for 1963–84, although promotional allowances, sampling distribution, and public entertainment expenditures must be subtracted from their figures for the years 1975–present to make the total advertising figures compatible with their earlier data. Data for 1955–62 are taken from Schmalensee [22]. The Schmalensee data and the FTC data overlap for the years 1963–67, and are nearly equal. For the years 1952–54, the Schmalensee data for 1955–67 were backcast using a set of exogenous variables for 1955–67 which were found to perform well: total reported advertising expenditures on selected media from various issues of *Advertising Age* [5], wholesale price, retail price, total quantity, and population 14 and older. Nominal expenditures were adjusted to 1980 dollars.

APPENDIX B

Simultaneous Estimation of the Model and Autocorrelation Coefficients

We estimate the autocorrelation coefficient and the model's coefficients simultaneously by adding the autocorrelation coefficient times the lagged error term to equation (4). Denote the right-hand side of equation (4) as the function $\psi(\alpha, X_t) + u_t$ where α is the vector of coefficients in (4) and X_t is the vector of predetermined variables in (4). The lagged predetermined variables are elements of X_{t-1}. With autocorrelation, $u_t = \rho u_{t-1} + \epsilon_t$, where ρ is the autocorrelation coefficient and ϵ_t is a normal error term. The lagged error term is $u_{t-1} = \{\ln Q_{t-1} - \psi(\alpha, X_{t-1})\}$, so $\rho u_{t-1} = \rho\{\ln Q_{t-1} - \psi(\alpha, X_{t-1})\}$ is added to equation (4) and NOLS is used to estimate ρ and α simultaneously.

In this correction for autocorrelation we are estimating only one more parameter than in the original model because we constrain the coefficients associated with lagged variables to equal the coefficients associated with the contemporary variables. The equation which is estimated in the autocorrelation correction is:

$$\begin{aligned}\ln Q_t = {} & \pi\alpha_0 + \pi\alpha_1 \ln P_t + \pi\alpha_2 \ln Y_t \\ & + \pi[\alpha_3 + \mu_1(D64_t) + \mu_2(D68_t) + \mu_3(D71_t) + \mu_4(D79_t)] \\ & \times \ln\{A_t + (1-\delta)A_{t-1} + (1-\delta)^2 A_{t-2} + \cdots + (1-\delta)^m A_{t-m}\} \\ & + (1-\pi)\ln Q_{t-1} \\ & + \rho\{\ln Q_{t-1} - (\pi\alpha_0 + \pi\alpha_1 \ln P_{t-1} + \pi\alpha_2 \ln Y_{t-1} \\ & + \pi[\alpha_3 + \mu_1(D64_{t-1}) + \mu_2(D68_{t-1}) + \mu_3(D71_{t-1}) + \mu_4(D79_{t-1})] \\ & \times \ln\{A_{t-1} + (1-\delta)A_{t-2} + (1-\delta)^2 A_{t-3} + \cdots + (1-\delta)^m A_{t-m-1}\} \\ & + (1-\pi)\ln Q_{t-2})\} \\ & + \epsilon_t\end{aligned}$$

where ϵ_t is the error term. This is the equation which well-known iterative methods approximate.

NOTES

1. Ashenfelter and Sullivan [1] test pricing behavior in the U.S. cigarette market. Their results support neither monopoly behavior *nor* competitive behavior [1, especially pp. 496–497]. This is consistent with the hypothesis of price leadership in the market (Scherer [21]).
2. The Federal Communication Commission's Fairness Doctrine of 1968 allowed one free broadcast anti-smoking warning for every four broadcast cigarette commercials. The broadcast advertising ban of 1971 eliminated the free anti-smoking warnings.
3. The model does not distinguish between increases in consumption due to new smokers and increases in consumption by individuals who were already smokers. Interesting studies which do are Jones [13] and Jones [14], which make use of detailed panel data available for the U.K. Such data are not available for the U.S.
4. Kao and Tremblay [15] suggest that the dependent variable might be expressed in terms of the amount of tobacco consumed, but their results indicate that the choice of dependent variable does not alter the estimates very much.
5. Durbin's $\hat{K}_2$ is the two-sided version of Durbin's h (Durbin, [6]). We use $\hat{K}_2$ because Durbin's h is a one-sided test for positive autocorrelation, whereas the results for the equations with 6, 8, and 10 lags suggest negative autocorrelation.

6. Results were also obtained for lags of 3, 5, 7, and 9, but these add little to the discussion and are excluded from the table.
7. Autocorrelation is not indicated in equations with 0, 1, 2, and 4 lags. The $\hat{K}_t$ statistics are 1.08 1.08, 0.193, 2.132, and 0.656.
8. Complete depreciation of advertising effects within one year is also consistent with Ashley, Granger, and Schmalensee [2]. Our equations with lagged advertising terms have unreasonable and/or statistically insignificant results associated with the coefficients for income and the 1964 Surgeon General's report.
9. The exogenous variables are chosen to obtain the results of a simultaneous supply and demand system in the spirit of Bishop and Yoo [4]. In P_t price is regressed on $\ln Y_t$, $\ln A_t$, $\ln Q_{t-1}$, ln (wages in the cigarette industry), and ln (cost of capital for cigarette firms).
10. α_1 is not significant in a two-tailed test. However, since a negative price elasticity is expected, a one-tailed test is appropriate. In addition, the one-tailed test is uniformly most powerful (UMP), whereas no UMP test exists for the two-tailed test [11]. In the analysis, *t* values for a nonlinear function f(x) of the vector x are found by calculating the asymptotic standard error $[(\partial f/\partial x)\Sigma(\partial f/\partial x)^T]^{1/2}$ where Σ is the relevant covariance matrix and T is the transpose operator.
11. This reflects the fact that D68 = 1 only for the years 1968–70.
12. One can think of the signals as being generated by the printed health warning on cigarette advertisements as required by the Cigarette Labeling and Advertising Act of 1965. Then the relationship $\sigma S_t = A_t$ holds for all periods when government policies were in place except for the initial year 1964. But since the first Surgeon General's report was published in 1964 (so that the warnings were fresh in people's minds) we will assume that the relationship holds for the entire period 1964–1984.
13. One might then wonder why the firms advertise when the signal elasticity overwhelms the advertising elasticity. A plausible explanation is that it is privately optimal: while advertising has the indirect effect of decreasing aggregate demand, it also has the effect of redistributing market share. So long as the gain to the firm offsets its share of the loss to the industry, the firm has an incentive to advertise.
14. The impact of the health warnings and media policies were substantial. Our regression results indicate that the effect of the 1964 health warning was equivalent to a tax increase sufficient to raise the 1964 price of a pack of cigarettes 219.4 percent. Similarly, the effects of the 1968 Fairness Doctrine, the 1971 broadcast advertising ban, and the 1979 health warning were equivalent to raising the price 126.2 percent, 175.7 percent, and 110.7 percent respectively. The details of these calculations are available from the authors upon request.
15. In addition to direct enforcement costs, Hamilton [9] suggests that advertising is the main competitive instrument in the cigarette industry so an advertising ban may result in even less competition in the industry. In fact, the industry recently began to compete in price through coupons distributed in the print media. These coupons are considered as advertising under the proposed policy and would be banned. In addition, disallowing advertising would discourage cigarette firms from developing safer cigarettes after the policy was implemented since they would not be permitted to convey the relative benefits of new (as yet undeveloped) cigarettes to consumers.
16. The problem was so severe in Hamilton [9] that he used price and income coefficients calculated by others, multiplied them times the price and income variables, and subtracted the results from his quantity variables. He then regressed these results on explanatory variables other than price and income.
17. Hamilton [9] uses one series which is longer than ours; but his series is based on corporations' tax deductions for advertising, while ours is based on actual advertising expenditures.

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