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 Cancer Society, 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 Balaghi and Levin [3], Hamilton [9], Schmalensee [22], and Smeeding et al. [23].

The minority view that advertising increases consumption is supported by Bishop and Yoo [4], Fujita [8], and Kao and Tremblay [15] for the U.S., McGuinness and Cowling [18], Randell [20], and Wirt and Pass [25] for the U.K.; Lee [16] for Switzerland, and McLeod [10] for Australia. But these studies have shortcomings. McGuinness and Cowling [18], Randell [20], and Wirt and Pass [25] use incomplete estimates of advertising expenditures. Lee [16] and McLeod [10] have no advertising data at all; they use advertising bans variables as proxies for advertising. Fujita [8] uses wholesale price rather than retail price. 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. 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.

Balaghi 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 dummy 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...
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 Radnor [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 (D68 = 1 for years 1968 through 1970, 0 otherwise), for the national 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 assure that the that the "desired" per capita consumption level Q* is given by

$$\ln Q^* = \alpha_0 + \sigma_1 \ln P_t - \sigma_2 \ln Y_t + \beta_1 \ln (D64_t) + \beta_2 \ln (D68_t) + \mu_1 \ln (D71_t) + \mu_2 \ln (D79_t)$$

where $$Q*$$ is in thousands of cigarettes, $$P$$ is the real retail price per thousand (including taxes), $$Y$$ is the real per capita disposable income, $$D46$$, $$D68$$, $$D71$$, and $$D79$$ are the dummy variables, and $$\sigma$$ is the per capita demand shift (which we call "goodwill") created by current and past advertising.

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

$$\ln Q_{t+1} - \ln Q_t = \psi (\ln Q^* - \ln Q_t)$$

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

(1) $$\ln Q_t = \alpha_0 + \sigma_1 \ln P_t + \beta_1 \ln (D64_t) + \beta_2 \ln (D68_t) + \mu_1 \ln (D71_t) + \mu_2 \ln (D79_t) + (1 - \psi) \ln Q_{t-1} + u_t$$

The coefficients are elasticities. For example, $$\psi$$ is the short-run price elasticity of demand while $$\sigma$$ is the long-run (full adjustment) price elasticity of demand. We assume that per capita goodwill changes as follows:

(2) $$\Delta G_t = G_{t-1} - G_t = A_t - B G_{t-1}$$

where $$A_t$$ is the difference operator, $$A_t$$ is real per capita advertising expenditures at time $$t$$, and $$B$$ is the depreciation rate for goodwill. Rearranging terms in equation (2) and substituting recursively for $$G_{t-2}, \ldots$$ yields

(3) $$G_t = A_t + (1 - B) A_{t-1} + (1 - B)^2 A_{t-2} + \cdots + (1 - B)^n A_{t-n} + \cdots$$

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

(4) $$\ln Q_t = \alpha_0 + \sigma_1 \ln P_t + \beta_1 \ln (D64_t) + \beta_2 \ln (D68_t) + \mu_1 \ln (D71_t) + \mu_2 \ln (D79_t)$$

$$\times \ln [A_t + (1 - B) A_{t-1} + (1 - B)^2 A_{t-2} + \cdots + (1 - B)^n A_{t-n} + \cdots + (1 - B)^m A_{t-m}] - \psi \ln Q_{t-m} + u_t$$

**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 tested a remittum term as in Baltagi and Levin [3], but the coefficient on the remittum term always took the wrong sign. For this reason, and because we find advertising to depreciate very quickly, we do not report these 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] $$R^2$$, $$\chi^2$$. 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 $$R^2$$, $$\chi^2 > 3.84$$ 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. The zero-lag case assumes $$m = 0$$, equivalently, $$\psi = 0$$. Autocorrelation seemed to be present in the equations with 4, 6, and 10 lags, so the correction procedure was applied in these regressions. The estimates and $$t$$ values of the autocorrelation coefficient $$\rho$$ 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 ($$\psi$$) 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.

We use Hausman's [10] test statistic to test price endogeneity in the equation with no lagged advertising terms. We calculate $$\omega = \beta \hat{B} - \hat{q}$$ in 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. Details regarding the instrument and the Hausman test are available from the authors upon request. $$\hat{p}$$ 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, $$\hat{p} > \chi^2(9)$$ with critical value 14.0837 at the 10 percent level. We calculate $$m = 1.4999$$. 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 ($$\sigma_1$$) is $$-10.1$$ with a t-ratio of $$-3.532$$. This is significant at the 5 percent level in a one-tailed test. Our estimate is smaller in absolute value than the estimate of $$-20.2$$ 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 $$-4.06$$ to $$-1.218$$. While our price elasticity is smaller in absolute value than these estimates, it supports the popular notion of a relatively high price-induced demand curve for cigarettes due to the habitual nature of cigarette consumption.

Our long-run price elasticity ($$\sigma_1$$) 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 $$-0.32$$ 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 ($$\psi$$) is nearly 0.7 and is close to the estimate of Schmalensee [22], which is 0.73.

Our short-run income elasticity is 0.29 with t-ratio 3.032 and is close to Schmalensee's [22] estimate.
## Table 1

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<th>Coefficients</th>
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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 $\xi$ implicitly equals 1.

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

of 0.32. Previous estimates range widely from -0.004 (Balltgi and Levin [1]) to 1.482 (Kao and Treadwell [15]). Our result supports the intuitively plausible notion that a change in income will have little impact on consumption.

$C_t = A_t$, when no lagged advertising terms is included. Our short-run advertising elasticity ($\mu_0$) is 0.00 with a t-ratio of 2.20 while the long-run advertising elasticity ($\mu_5$) is 0.20 with a ratio of 2.62. 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 $\mu_5$ is the complete advertising ban. This would 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 $\tau(n, \mu_1, \mu_2, \mu_3, \mu_4, \mu_5, \mu_6, \mu_7)$, where $\omega(n, \mu_1, \mu_2, \mu_3, \mu_4, \mu_5, \mu_6, \mu_7)$. The estimates (1 values) are
advertising figures compatible with their earlier data. For 1955-67 are taken from Schmalensee [22]. The Schmalensee data and the FTC data overlap for the years 1953-67, 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 \( \sigma(X_t) + \alpha_u + \varepsilon \), where \( \alpha_u \) 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 \( X_t \). With autocorrelation, \( \alpha_u = \alpha_u + \varepsilon \), where \( \varepsilon \) is the autocorrelation coefficient and \( \varepsilon \) is a normal error term. The lagged error term is \( \varepsilon = [\varepsilon - (X_{t-1}) + \varepsilon = \alpha (X_{t-1}) ] \), is added to equation (4) and GOLS is used to estimate \( \varepsilon \) and \( \alpha \) simultaneously.

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

\[
Y_t = \beta_0 + \beta_1 Y_{t-1} + \beta_2 X_{t-1} + \cdots + \beta_t X_{t-t} + \varepsilon_t,
\]

where \( \varepsilon_t \) is the error term. This is the equation which well-known iterative methods approximate.

**NOTES**

1. Ashenfelter, O'ley and Sullivan [1] test pricing behavior in the U.S. cigarette market. Their results support neither monopoly behavior nor competitive behavior [1], especially pp. 490-497. 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 five 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. The smoking studies which are done 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 Trentham [17] 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 K, is the two-sided version of Durbin's (Durbin, [6]. We use \( \hat{K} \) because Denby's \( K \) 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 lag 5, 3, 7, and 9, but these add little to the discussion and are excluded from the tables.

7. Autocorrelation is not indicated in equations with 6, 1, 2, and 10 lags. The K estimates are 1.081.100, 0.013, 2.132, and 0.056.

8. Complex depreciation of advertising effects within one year is also consistent with Ashby, Grainger, and Schmalensee [2]. Our equations with lagged advertising terms have unreasonable and/or statistically insignifi-

9. The exogenous variables are chosen to obtain the results of a simultaneous supply and demand system in the

10. The results in this test is uniformly more powerful (UMP), whereas no UMP test exists for the two-tailed test [12]. In the analysis, \( \alpha \) values for a function \( f(x) \) of the vector x are found by calculating the asymptotic standard error (ASE) of \( f(x) \) to \( \alpha \) where \( \alpha \) is the relevant observation minus T and \( \alpha \) is the transpose operation.

11. The results that \( D_{67} = 0 \) only for the years 1960-70.

12. One can think of the signals as being generated by the printed health warnings on cigarette advertisements as required by the Cigarette Labelling and Advertising Act of 1965. Then the relationship \( \alpha \) = 0 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 electricity outweighs 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 nationwide. Our regression results indicates 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 214.4 percent. Similarly, the effects of the 1964 Fairness Doctrine, the 1971 broadcast advertising ban, and the 1979 health warning were raising the price 128.2 percent, 175.7 percent, and 113.7 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 an 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, deleting 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 we unsure 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 larger than ours, but his series is based on corporations' tax deductions for advertising, while ours is based on actual advertising expenditures.

**REFERENCES**


A Note on the Bounded Solution of the Bilateral Monopoly Model

Chin W. Yang

Consider a case of bilateral monopoly in which an intermediate product (e.g., tobacco leaves) is sold to a single buyer who uses it to produce final product (e.g., cigarettes). The buyer of the intermediate product sells it in a competitive market at a constant price $P_2$ and a single seller of the intermediate product purchases raw input (or a combined input) $x$ from a competitive market at a constant price $r$ in order to produce $y$. These production functions are shown below:

\[ f: X \rightarrow Y_1 \]

\[ g: Y_2 \rightarrow Z \]

where $f(x) = y$ is a functional mapping from input space $X$ into intermediate output space $Y_1$ such that $f(x) = y$, then $x = x_0$. This is to say that the inverse production function, $x = f^{-1}(y)$, exists uniquely for the upstream monopolist; $g(y) = z$ is a functional mapping from $Y_2$ into the final output space $Z$ and $X$, $Y_1$, $Y_2$, and $Z$ are sets of nonnegative real numbers $\mathbb{R}^+$. Note that we are only concerned with the interior of the intermediate product set $Y = Y_1 \cap Y_2$ so that the chain rule applies.

It is commonly assumed that neither the single buyer nor the single seller could dominate the market of the intermediate product $y$. As a consequence, there may well be a tendency to move toward an "optimum" quantity of $y$ that maximizes their joint profits.

Maximize $x_1 + x_2 = P_1 y + P_2 y + P_3 y - r x$

$y \in Y, P_2 \in \mathbb{R}^+$

\[ P_2: g(y) = P_3 y + P_3 y - r f(y) \]

\[ \text{[3]} \]