excitable properties of the heart may be of value in preventing sudden cardiac death, which usually results from ventricular fibrillation.

NEAL A. SCOTT, REGIS A. DESILVA
BERNARD LOWN
RICHARD J. WURTMAN
Cardiovascular Research Laboratories, Department of Nutrition, Harvard School of Public Health, Boston, Massachusetts 02115 and *Laboratory of Neuroendocrine Regulation, Department of Nutrition and Food Science, Massachusetts Institute of Technology, Cambridge 02139*

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Cigarette Smoking in the 1970's: The Impact of the Antismoking Campaign on Consumption

Abstract. *Per capita cigarette consumption has declined annually since 1973. This downturn represents a new nonsmoking ethos, which is also reflected in the legislative successes of the nonsmokers' rights movement. Time series regression analysis of cigarette demand suggests that, in the absence of the antismoking campaign, consumption would have exceeded its 1978 level by more than a third.*

Adult per capita cigarette consumption in the United States rose throughout the century until 1964, the year of the Surgeon General's report on smoking and health (*1-3*) (Fig. 1). Continued increases would have been expected in the absence of an antismoking campaign, especially with more women smoking (*4*) and decreases in the nicotine content of cigarettes. According to the nicotine regulation hypothesis, the latter should have increased the typical smoker's cigarette consumption (*5*). Yet per capita consumption fell in 1968 through 1970; this was the first time that it had dropped for more than two consecutive years. Although causality has not been conclusively established, numerous investigators have attributed the decreases to antismoking messages carried on television and radio as a result of the Federal Communication Commission's Fairness Doctrine. Several downturns occurred earlier in years with specific major antismoking "events" such as the first attention in the popular press in 1953 and 1954 to the link between smoking and illness and the Surgeon General's 1964 report (*3, 6*).

Since 1973 per capita cigarette consumption has fallen approximately 1 percent per year. This downturn cannot be associated with discrete antismoking events, but it has occurred during a decade in which an active nonsmokers' rights movement has developed.

To assess the effect of the antismoking campaign on cigarette consumption and explore the relation between the nonsmokers' rights movement and the decreases in consumption, I have revised my earlier time series analysis of adult per capita demand for cigarettes (*7*). As in other studies (*6*), I account for the effects of smoking-and-health publicity by inclusion in a demand regression equation of binary (dummy) variables with a value of 1 in years of adverse publicity and 0 in other years. In order to examine the decline in consumption since 1973, I include a measure of the effectiveness of the nonsmokers' rights movement, namely, the percentage in a given year of the adult population residing in states with laws that restrict smoking in public places.

A least-squares regression for the years 1947 to 1978 yields the demand equation (*8*)

$$\begin{aligned} C_t = {} & -4621.2 - 16.36^*P_t + 0.6896^*C_{t-1} + 1824.7^*Y_t - 131.39^{**}D53_t - 305.41^*D54_t - \\ & 195.25^*D64_t - 110.05^{***}D68_t - 226.72^*D69_t - 79.3D70_t - 837.81^*L_t \\ & [^*P < .01,\ ^{**}P < .05,\ ^{***}P < .10; \\ & F(10, 21) = 76.56;\ R^2 = .9733] \end{aligned}$$

where C_t is adult per capita cigarette consumption in year t (mean = 4055.8); C_{t-1} is lagged consumption, which captures the effect of habit; P_t is relative real cigarette price in year t (index with $P_{1967} = 100$) and assesses price responsiveness of consumption; Y_t is the natural logarithm of the last two digits of year t and reflects increases in the smoking population due to the diffusion of the behavior, particularly among women, and increases in smokers' consumption levels (*9*); $D53_t$ and so forth are the dummy variables for 1953, 1964, 1968, 1969, and 1970, which provide estimates of the impact on consumption of antismoking activities during these years; $D54_t$ is 0 before 1954 and $0.5^{(t - 1954)}$ in 1954 and on, a dummy for the second year of smoking and health publicity, which equals 1 in 1954 and has a continuing, though rapidly diminishing, additional effect that reflects additional publicity through the mid-1950's; and L_t is the state law population measure in year t.

The habit factor C_{t-1} indicates the persistence of close to 70 percent of the preceding year's per capita consumption, with other factors held constant. The P_t coefficient implies a price elasticity of demand of -0.37 at the means of dependent and independent variables. Estimates of elasticity range from close to 0 to -1, with a consensus estimate of -0.45 to -0.5 (*10*). However, some investigators suggest that either elasticity has been falling or it has been systematically overestimated (*11, 12*).

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With the exception of $D70_t$ (P = .2), the dummy variables imply that significant decreases in per capita consumption occurred in each year of prominent adverse smoking-and-health publicity; the decreases range from about 100 to 300 cigarettes, or 2.5 to 7.7 percent of annual per capita consumption. Although each dummy has a value of 1 only in the year of adverse publicity (except $D54_t$), inclusion of lagged consumption in the equation means that publicity-induced decreases in consumption in year t persist in ensuing years, though at a diminishing rate. For example, publicity surrounding release of the Surgeon General's report in 1964 was associated with a decrease in consumption of 195.25 cigarettes in 1964—the coefficient of $D64$. When this decrease is multiplied by the coefficient of the habit factor (0.6896), it appears that publicity is associated with a decrease in consumption of 134.64 cigarettes in 1965, and 92.85 cigarettes in 1966 (that is, 134.64 × 0.6896), and so on.

The index of the success of the nonsmokers' rights movement (L_t) is strongly correlated with the decreases in consumption dating from 1973, both in statistical significance (P < .0001) and absolute magnitude: L_t rises from 0.011 in 1964 to 0.723 in 1978; when these values are multiplied by the coefficient −837.81, the associated new decreases in annual consumption are 9 cigarettes in 1964 and 606 in 1978.

Although the restrictions placed on smoking by new laws might force small reductions in cigarette consumption, it seems implausible that physical restraints could account for the magnitude of the effect suggested by the regression. An alternative explanation is that the effect on behavior of the antismoking campaign lagged behind changes in knowledge and attitude (*13*). The passage of the laws, rather than shaping opinion and affecting behavior, may reflect opinion and coincident behavior change.

The decreases in cigarette consumption of the 1970's occurred during a period in which certain conditions favored increases; for example, relative to the general rise in consumer prices, cigarette prices dropped annually after 1972. One measure of the effectiveness of the antismoking campaign is a comparison of actual cigarette consumption with that which would have been expected in the absence of the campaign. The regression equation permits estimation of the latter. This is plotted as the dashed line in Fig. 1. For any given year, the new and residual decreases associated with the specific antismoking events (the publicity represented by the dummy variables and L_t) can be added to actual consumption to make the estimate. The difference between estimated and actual consumption, divided by actual consumption and multiplied by 100 percent, is the percentage by which consumption would have exceeded its realized value had the antismoking campaign not materialized. In 1978, this was 36.4 percent. In Fig. 1 this difference reflects the vertical gap between the endpoints of the solid line, representing actual consumption, and the dashed line, representing anticipated consumption.

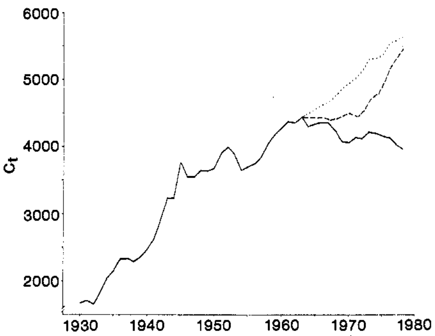

Fig. 1. Per capita cigarette consumption (C_t): the solid line represents actual consumption; the dotted line is consumption predicted in absence of an antismoking campaign and with an assumed price constancy; the dashed line is consumption predicted in absence of a campaign but with actual prices experienced.

The flatness of the anticipated consumption line from 1964 through 1972 reflects the restraining influence of continuous relative price increases resulting primarily from state and local excise taxes; P_t rose from 92.70 in 1963 to a high of 107.98 in 1972 and then began to drop, reaching 92.02 in 1978. It seems reasonable to attribute the large growth in taxation through 1972 to the antismoking campaign (*3, 14*). By assuming that the relative cigarette price would not have changed in the absence of the campaign, I can estimate the levels that per capita consumption would have reached in the absence of both publicity alone and publicity-induced price effects (Fig. 1). The combined effect produces the consumption path plotted as the dotted line in Fig. 1. This measure shows that in 1978 consumption would have been 41.5 percent greater than that which actually occurred. In any year, the effect on consumption of publicity only is the distance between the dashed and solid lines in Fig. 1; the distance between the dotted and dashed lines represents the effect on consumption of price changes induced by the antismoking campaign.

Alternative specifications of the demand function and means of estimating the impact of the antismoking campaign on consumption produce slightly different quantitative results, but the qualitative findings are unaffected (*15*). (i) In the absence of the antismoking campaign, cigarette consumption would have been considerably higher than it is today (*16*). (ii) The major impact on consumption of publicizing smoking-and-health information has lagged behind knowledge and attitudinal changes. By 1971, the midpoint of the period studied, the impact on consumption of publicity alone was less than 10 percent of actual consumption. By 1978, it was three to four times as great. (iii) The decreases in cigarette consumption since 1973 correlate highly with the legislative successes of the nonsmokers' rights movement. This correlation seems unlikely to reflect causation. Rather, both declining consumption and growth in legislation probably reflect a prevailing nonsmoking ethos and the conversion of modified knowledge and attitudes into behavioral change.

KENNETH E. WARNER

School of Public Health,
University of Michigan, Ann Arbor 48109

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7. Variables excluded from the analysis are discussed in (*3*).
8. Autocorrelation was tested with the Durbin H, a variant on the Durbin-Watson test designed for the case of inclusion of a lagged dependent variable. Scatter plots revealed no heteroscedasticity.
9. The Y_t can be criticized as implying continuous diffusion of smoking forever, albeit at a diminishing rate. However, the implied annual addition to per capita consumption is only 23 cigarettes in 1980, dropping to 18 by the year 2000. Including the effects through the lagged dependent variable, the suggested increase through the year 2000 is consistent with, even conservative compared to, diffusion in earlier periods and diffusion which might have been antici-

pated in the absence of the smoking and illness link, especially with the increase in smoking by women. At some point diffusion per se would have ceased. Thus this model should not be interpreted as applying to the 21st century.
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14. The curtailment of excise tax growth partly reflects concerns about bootlegging (*12*).
15. The method used differs from and produced slightly more conservative estimates than that in my earlier study (*3*). Still more conservative estimates result if a logistic specification is substituted for the linear model.
16. Other behavioral changes are of much greater magnitude, and possibly health significance, than decreases in the numbers of cigarettes smoked. For example, sales-weighted tar and nicotine per cigarette have fallen by 50 percent since the 1950's. While the health implications of this change are unestablished, there is histologic evidence that the change may result in significantly decreased damage to health [O. Auerbach, E. Hammond, L. Garfinkel, *N. Engl. J. Med.* **300**, 381 (1979)].
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Solar Wind Data and Ionospheric Potential

In their article on "Solar wind control of the earth's electric field" Markson and Muir (*1*) correctly indicate that Barouch and Burlaga (*2*) associated rapid increases in solar wind velocity with decreases of the galactic cosmic-ray intensity. However, we did not claim to find any association between the value of the solar wind velocity and the cosmic-ray intensity. Indeed, our figure 8 (*2*) shows that, for the period we considered, this association could not easily be established. Our interpretation was that magnetic blobs are formed at the head of interplanetary streams, where the faster plasma catches up with the slower, earlier ejected, plasma.

If a relation between ionospheric electricity and cosmic-ray ionization of the upper atmosphere is sought, it seems that a direct approach should be investigated. Several cosmic-ray groups routinely fly balloons directly measuring the intensity of secondary particles as a function of altitude; this is probably the most straightforward way. Failing access to these data, one could look at the hourly values of the neutron monitor cosmic-ray detectors, or of cosmic-ray detectors aboard earth-orbiting satellites, which are published monthly in *Solar Geophysical Data* and are available on request from the World Data Centers. The indirect and somewhat unreliable associations between value of solar wind velocity and solar wind velocity changes, solar wind velocity changes and magnetic field enhancements (blobs), blobs and Forbush decreases and, Forbush decreases and upper atmospheric ionization form a chain where many unexpected other phenomena can distort the meaning of the numerical correlation that can be calculated.

A quick check of the daily averaged data presented by Markson and Muir in their table 1 is instructive in this respect [this time scale is more appropriate than hourly values, since the correlation time of the solar wind velocity is about 3 days (*3, 4*)]. These values were divided into three classes: those approximately coincident in time with the head of an interplanetary stream, those approximately coincident in time with the tail of an interplanetary stream, and those where the interplanetary data were too scarce or complex to be classified.

For the first two classes (six and seven points, respectively) the averages of the potentials given by Markson and Muir are 197 ± 14 and 224 ± 23 kV. Clearly, the statistics are as yet insufficient for definitive conclusions to be drawn from these data, although the ratio between these values is in the direction predicted by Markson and Muir.

E. BAROUCH
Centre de Recherches en Physique de l'Environnement Terrestre et Planetaire, Issy-les-Moulineaux 92131, France

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The suggestion for a direct comparison between cosmic radiation data and ionospheric potential (V_I) variations is appropriate since in the referenced article (*1*) this relationship was deduced from inverse correlations between (i) solar wind velocity and V_I and (ii) solar wind velocity and cosmic radiation. I recently completed a study of cosmic rays and V_I and found the expected direct correlation. (Ionospheric potential is a measure of the earth's overall fair-weather electric field intensity.) The slope of the regression line is such that a 10 percent change in cosmic-ray flux measured by a neutron monitor corresponds to an approximately 15 percent change in the earth's electric field intensity.

This analysis was not performed earlier because not all the desired data had been acquired and we did not wish to delay publication of the finding that solar wind velocity controlled a fundamental property of the upper and lower atmosphere. In statistical studies of sun-weather relations, use of the solar wind velocity, a continuous property of the interplanetary medium, offers a considerable advantage over use of crossings of solar magnetic sectors as timing marks; the latter are nonparametric events which occur only two to four times a month.

A comprehensive examination of cosmic radiation data and V_I variations has now been completed through analysis of two independent V_I data sets: (i) the Bahamas measurements used in the solar wind analysis (*1*) and (ii) a recently acquired listing of German soundings provided by R. Mühleisen and H.-J. Fischer (Astronomisches Institut der Universität Tübingen). The Bahamas aircraft soundings (N = 92 for the purpose of analysis) were made during a 3-month period (December 1971 to March 1972), and the German balloon data (N = 237 for analysis) were obtained over a 17-year period (1959 to 1976). These are the only V_I data sets that are sufficiently extensive and coherent for the required analysis.

Balloon measurements of stratospheric ionization have not been conducted with sufficient frequency to allow meaningful comparison with V_I soundings, which have also been obtained sporadically. However, as noted by Barouch, the required unbroken time series of cosmic radiation measurements from satellites and neutron monitors exist; these have been correlated with both sets of V_I data. The satellite measurements used are from IMP-4 during the period June 1969 to December 1972; they provide the longest coherent satellite data set available covering the period of the Bahamas records. The Mount Washington neutron

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