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Willingness to Pay for Ozone Control: Inferences from the Demand for Medical Care[1]

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The paper applies a discrete choice version of the household production approach to the valuation of nonmarket goods. Willingness to pay for tropospheric ozone control is estimated using medical care demand under assumptions of input necessity and weak complementarity. In example calculations, individuals living in high ozone areas are willing to pay over \$170 annually for an environment in which ozone concentrations never exceed 12 pphm. Willingness to pay figures are two to four times larger than medical expense savings caused by the same ozone reductions. Estimates obtained are compared with results of previous studies, and proposed ozone control measures are discussed. © 1991 Academic Press, Inc.

1. INTRODUCTION

Estimating monetary benefits of increased nonmarket commodity supplies has proved to be a vexing although important aspect of environmental policy formulation. Several methods based on hedonic prices, travel costs, contingent values, and direct monetary damages have been developed to estimate willingness to pay; yet, due both to theoretical reasons and to data availability, no single method has come close to winning universal approval. Even the hedonic price method, used by some investigators (e.g., Brookshire *et al.* [1]) as a standard of comparison by which to evaluate the efficacy of other methods, increasingly has been called into question. Basic issues in identifying key demand and supply parameters recently have received extensive discussion by Brown and Rosen [2], Bartik and Smith [3], Bartik [4], and McConnell and Phipps [5], and a good survey of issues raised can be found in Mendelsohn [6]. Also, Atkinson and Crocker [7] and Graves *et al.* [8] have demonstrated that a single hedonic regression can produce an uncomfortably large range of willingness to pay estimates depending on how it is specified.

In light of these and other problems with commonly used benefit estimation methods, it may be worthwhile to consider an alternative procedure that makes use of a household production function framework and infers demand for a nonmarket

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commodity from demand for a complementary private good. Conceptual aspects of this approach have been treated by Hori [9] and Bockstael and McConnell [10]; however, it seldom has been applied because relevant data often are unavailable and because of difficulties in establishing whether certain theoretical conditions are met.

This paper provides an illustration of the household production approach, although it does not completely solve all associated methodological problems. In particular, conservative willingness to pay estimates for tropospheric ozone control are inferred from the demand for medical care. Example calculations show that individuals living in urban areas where the 1-hour peak ozone concentration exceeds the current federal standard of 12 pphm on 117 days, with an average concentration on those days of 18 pphm and a high of 35 pphm, are willing to pay over $170 annually for an environment in which peak ozone concentrations never rise above 12 pphm. Willingness to pay figures are approximately two to four times larger than reductions in medical expenses that would result from the same reductions in 1-hour peak ambient ozone levels. Section 2 outlines a simple household production model. Section 3 describes data obtained from a panel study of Southern California residents designed specifically to allow the estimation of health relationships. Section 4 presents empirical results based on probit estimates of a medical care demand equation. Implications and conclusions are drawn out in Section 5 by comparing estimates of willingness to pay presented here with related estimates obtained in previous studies and by discussing ozone control measures recently proposed at the federal level and in California.

2. MODEL

A one-period model specifies utility (U) as a function of market goods (X), health (H), and exposure to air pollution (α). For simplicity, X is treated as a composite good and H is treated as a nonnegative index of health attributes. Health is home produced by combining medical care received during the current period ($M \geq 0$) with medical treatment and health related information acquired in previous periods ($\overline{M} \geq 0$), genetic capital endowments ($K \geq 0$), and exposure to air pollution. This specification implies that people can build up a stock of information through contacts with the health care delivery system that makes them better at treating themselves in related settings as well as more knowledgeable consumers of medical services. Health related information is assumed never to be forgotten, so the time at which it was acquired is irrelevant. Modeling health decisions in a household production framework has been utilized elsewhere in the literature (e.g., Grossman [11], Rosenzweig and Schultz [12, 13], Gerking and Stanley [14], Harrington and Portney [15]), where medical care is a private good input that often is treated explicitly.

In this paper, the decision to seek medical care in the current period (in order to obtain treatment and/or health related information) is framed in a discrete choice context. The health production function is written as

$$H_1 = H_1(M; \overline{M}, K, \alpha) \geq 0 \qquad \text{if } M > 0 \tag{1a}$$

$$H_0 = H_0(0; \overline{M}, K, \alpha)\begin{cases} = 0 & \text{if } M^* > \overline{M} \\ \geq 0 & \text{if } \overline{M} \geq M^*, \end{cases} \tag{1b}$$

where $H = 0$ denotes a state called poor health and $H > 0$ is an index reflecting varying states of good health, and where $M^* > 0$ is the critical stock of health related information required to maintain good health in the current period.[2] M^*, in turn, is determined by the extent of exogenous environmental insults experienced by the individual net of an individual's resistance to disease (determined by K). As shown, current medical care, M, is an essential input in the production of good health if the individual's previously acquired knowledge is inadequate to maintain good health. A person who faced a medical crisis (represented by a large M^*) might find his life threatened without diagnosis and treatment by a physician during the current period, while a person facing minor or familiar health problems (small M^*) could maintain good health using his stock of health related information as a substitute for current medical attention.[3] Of course, poor health could occur no matter how much past and present medical care has been obtained if the endowment of genetic capital is small enough or the extent of environmental insults is large enough. As demonstrated by Bockstael and McConnell [10], treating medical care as an essential input is one of two sufficient conditions for using its demand curve as a basis for welfare measurement. If in reality medical care is not an essential input in producing good health, changes in area behind the demand curve turn out to be a lower bound on true compensating variation for air quality changes (see Just *et al.* [16, Chap. 4]).[4]

The utility function, conditional on the discrete choice of whether to consume medical care during the current period, is

$$U_1 = U_1(X, H_1, \alpha) \tag{2a}$$

$$U_0 = U_0(X, H_0, \alpha). \tag{2b}$$

Weak complementarity between H and α ($\partial U/\partial \alpha = 0$ when $H = 0$), a second sufficient condition extensively discussed by Bockstael and McConnell [10], also is assumed to hold. This condition would hold trivially in the case where α does not enter the utility function, a specification often maintained in household production models of air pollution and health [14, 15, 17–19]. In the situation at hand, where α is an argument of the utility function, weak complementarity implies that an individual in poor health is indifferent to changes in air pollution levels. This

[2]This formulation of health production is inadequate for modeling lifetime decision making because, for example, it does not explicitly allow for investment in health capital, transition from good to poor health or death, or present consumption of medical care to affect future health. Consequently, the model is better applied to comparatively short observational periods in which large changes in H are unlikely. See Cropper [31] for discussion of health investment issues.

[3]M^* is the value of $\overline{M}$ which solves $\overline{H}_0 - H_0(O; \overline{M}, K, \alpha) \equiv 0$ for an arbitrarily small $\overline{H} > 0$. Several cases can be identified, including: (a) if M is essential, no real valued solution will exist; (b) if $M^* = 0$ is a solution, the M and $\overline{M}$ are inessential both individually and jointly; and (c) if there is a unique solution $M^* > 0$, then M and $\overline{M}$ are jointly essential in the sense that good health cannot be produced in the absence of both. If in addition $H_0(\cdot)$ has continuous first partials and $\partial H_0/\partial \overline{M} \neq 0$, then M^* is an implicit function of K, α, and $\overline{H}_0$, as assumed in the model.

[4]Note that the model could be respecified to allow all or at least certain types of medical care to be home produced, thus incorporating more fully the idea that most people do not consult physicians for every health problem faced. This extension is not pursued, however, because it would not alter the nature of the essential input problem. Medical care rendered by physicians still would have to be an essential input in the production of health or another input essential in home producing medical services would have to be identified.

condition means that when $H = 0$, a reduction in air pollution will not lead to health improvements, nor to any direct effect on utility (as might occur with an improvement in visibility).

While weak complementarity is an assumption about parameters of the utility functions in Eqs. (2), the estimation approach discussed in Section 4 recovers only parameters of a utility difference rather than parameters of the original utility functions. As a result, the empirical methods used here do not test this assumption. If the weak complementarity assumption is false, the compensating variation will be underestimated.

The full income budget constraint, conditional on the value of M, is

$$wT = q_X X + q_M M + wG(H_1) \tag{3a}$$

$$wT = q_X X + wG(H_0), \tag{3b}$$

where w denotes the wage rate, T denotes total time available, q_X and q_M denote the full, time inclusive prices of X and M, and $G(H)$ expresses time lost from market and nonmarket activities as a function of H, $G_H < 0$. Using the above equations, and following Small and Rosen [20], two conditional indirect utility functions are defined, giving maximum utility attainable depending on the choice of whether to seek medical care in the current period:

$$V_1(w, T, q_X, q_M, \overline{M}, K, \alpha) = \max\,(2\text{a}) \quad \text{subject to (1a) and (3a)} \tag{4a}$$

$$V_0(w, T, q_X, \overline{M}, K, \alpha) = \max\,(2\text{b}) \quad \text{subject to (1b) and (3b)} \tag{4b}$$

Thus, medical care is obtained if the utility difference $V_1(\cdot) - V_0(\cdot) > 0$ and a family of medical demand curves can be defined conditional on the value taken by $\overline{M}$, with the values of all other parameters held constant. Ideally, welfare effects of a change in α would be evaluated by setting $\overline{M} \leq M^*$ (so that medical care in the current period is an essential input) and measuring the change in consumer surplus behind the demand curve for M.[5]

3. DATA

Data were obtained from a sample of 226 residents of two Los Angeles area communities: 151 respondents lived in Glendora (a community with high oxidant air pollution levels) while 75 lived in Burbank (a community with oxidant pollution levels more like other urbanized areas in the United States but with comparatively high levels of carbon monoxide). All respondents were either nonsmokers or former smokers who had not smoked in at least two years, and all were household heads with full-time jobs (defined as at least 1600 hours of work annually).

The sample was drawn from participants in a prior study of chronic obstructive respiratory disease (Detels *et al.* [21, 22]) and included a disproportionate number of persons with compromised respiratory function. Seventy-six persons suffered from physician diagnosed breathing disorders and 50 persons suffered from self-reported chronic cough or chronic shortness of breath, while the remaining 100

[5] M^* does not appear as a parameter in the indirect utility function because all the information conveyed by M^* is already captured in $H_0(\cdot)$.

persons reported uncompromised respiratory function. Differences in medical demand and willingness to pay estimates between respondents with and without chronic lung disease are discussed in Sections 4 and 5.

Professionally trained interviewers contacted respondents several times over a 17-month period beginning in July 1985, but the last contact involved a much revised survey that did not focus on medical care. As a consequence, data used in the empirical work were collected during the first 12 months of the study. The first contact involved administration of an extensive baseline questionnaire in the respondent's home. Subsequent interviews were conducted by telephone. Including the baseline interview, the number of contacts with each respondent that yielded data on medical care varied from 2 to 5 with an average number of contacts of just over 4. Of the 928 total contacts, 201 were with respondents who reported physician diagnosed chronic lung disease and the remaining 727 were with respondents who did not.

Initial baseline interviews measured three groups of variables: (1) long-term health status, (2) contacts with the medical care delivery system, and (3) socioeconomic/demographic and work environment characteristics. Telephone follow-up surveys inquired further about medical care contacts.

Long-term health status was measured in two ways. First, respondents indicated whether a physician ever had diagnosed asthma (ASTHMA), chronic bronchitis (BRONCH), or another chronic respiratory disease such as emphysema, tuberculosis, or lung cancer. The dummy variable CHRONIC indicates whether a respondent reported any of the above physician diagnosed chronic lung diseases. Second, they stated whether they experience chronic shortness of breath and wheezing (SHRTWHZ) and/or regularly cough up phlegm, sputum, or mucous (FLEMCO) and whether they suffer from hayfever (HAYFEV).

Both baseline and follow-up surveys asked whether medical care, defined as a visit to a doctor's office, emergency care facility, or hospital, had been obtained during the two days preceding the survey. The survey did not ascertain whether an appointment had been scheduled during the two-day period. The binary dependent variable MED takes the value of unity if medical care was obtained and zero otherwise; medical care was obtained in 71 of the observations. A theoretically preferable variable might be the quantity of medical services consumed or amount of health related information received, but that quantity is difficult to measure and may be determined by the supplier. Variable MED is measurable and is a choice variable, indicating a willingness to enter the market for medical care. Neither the stock of health related information accumulated through prior contacts with the medical care delivery system ($\bar{M}$) nor the critical stock (M^*), however, could be directly measured in the survey. As discussed in Section 4, the panel structure of the data is used to determine effects of the stock of health information.

The two types of surveys also asked about the cost respondents incur when seeking medical care. A series of questions in the baseline survey began by asking whether respondents had a regular doctor (DOCREG). Those who answered in the affirmative then were asked about the typical out-of-pocket (net of insurance or other reimbursement) expense incurred for a visit to their doctor (PMED), as well as the commuting and waiting time required to see their doctor (TMED). The full price of medical care was computed as FPMED = PMED + WAGE $*$ TMED, where WAGE represents the respondent's hourly wage. For the small number of respondents who had no regular doctor, FPMED was computed using sample means of PMED and TMED together with the respondent's own WAGE. The sensitivity of estimated air pollution control benefits to using WAGE to measure the marginal value of time is investigated in Section 5.

If medical care was obtained in the two days prior to the interview, respondents also were asked their out-of-pocket expense and the time spent commuting and waiting for that particular visit. These measures are, of course, unobservable for the large proportion of respondents who did not obtain medical care. Moreover, a large proportion of the respondents who reported obtaining medical care did not report their out-of-pocket expense, possibly because of uncertainty about the extent to which the care obtained would be covered by insurance. As a consequence, the variable FPMED rather than the costs of the most recent doctor's visit is used as the price variable. FPMED may be a superior variable in any case, as the cost of medical care may not be known in advance, in which case FPMED may proxy for the expected cost of care.

In addition to WAGE, socioeconomic/demographic variables measured whether the respondent lived in Burbank or Glendora (BURB), as well as years of age (AGE), years of education completed (EDGRADE), occupation (BLUE = 1 if blue collar occupation), gender, race, and marital status. Also, respondents were asked whether they were exposed to toxic fumes or dust while at work (EXPWORK).

Finally, each contact with a respondent was matched to daily measures of ambient air pollution concentrations, humidity, and temperature for that day. As argued by Murdoch and Thayer [23], day-by-day data are superior to temporal means of these variables because they incorporate information on the probability distribution of outcomes. Air monitoring stations used are those nearest to residences of respondents in each of the two communities. Measures of the six criteria pollutants for which natural ambient air quality standards have been established were obtained: carbon monoxide (CO), nitrogen dioxide (NO2), ozone (O3), sulfur dioxide (SO2), lead, and total suspended particulates. Readings for lead and particulates, however, were available for only about 10% of the days during the survey period, forcing exclusion of those two pollutants from empirical work. Each of the remaining four pollutants were measured as maximum daily one-hour ambient concentrations. Maxima are used because epidemiological and medical evidence suggests that acute health problems of the type likely to induce a visit to the doctor may be more closely related to peak than to average concentrations. Since the survey did not record the day on which medical care was obtained, the air pollution variables entered are averages of one-hour maxima on the two days prior to the interview, to conform with the two-day measurement of medical demand. Daily high temperature (TEMP) and low relative humidity (HUMID) data similarly were averaged across the two days.[6]

4. EMPIRICAL RESULTS

As discussed in Section 2, the decision to seek medical care is a discrete choice dependent on the difference in indirect utilities with and without medical care.

[6]Measurement of medical demand on a two-day basis complicates the construction of pollution variables which are temporally consistent with the dependent variable. Other treatments of pollution data are possible, but none resolve the issue of the measurement of MED.

The utility difference is specified econometrically as

$$V_{1it}(\cdot) - V_{0it}(\cdot) = X_{it}'\beta + u_i + v_{it}, \tag{5}$$

where i and t index respondents and time, X_{it} is a vector whose first element is unity and whose remaining elements measure arguments of V_1 and V_0, β is a parameter vector, and u_i and v_{it} are random error components. The permanent components u_i capture unmeasured, individual-specific influences affecting whether medical care is sought, such as the stock of health related information acquired through prior contacts with the medical care delivery system ($\overline{M}$). The transitory components v_{it} capture random effects which vary both over individuals and over time during the sample period, including environmental insults not captured by α. The probability of obtaining medical care, conditional on u_i, is

$$\Pr(M_{it} = 1) = F_{it} = F(X_{it}'\beta + u_i),$$

where $M_{it} = 1$ if medical care is obtained and 0 otherwise, and $F(\cdot)$ is the symmetric distribution of V_{it} conditional on u_i. When the conditioning on u_i is removed and all observations are considered, the sample log-likelihood function is

$$\ln L = \sum_i \ln \int \pi_t F_{it}^{M_{it}} (1 - F_{it})^{1-M_{it}} g(u)\, du, \tag{6}$$

where $g(\cdot)$ is the marginal density of the u_i. Assuming that both u_i and v_{it} are normally distributed yields the random effects probit model. Probit was chosen over logit because, as discussed by Maddala [24], the logistic distribution severely restricts the error correlations in a random effects model. Also, a random effects approach was chosen because there does not appear to be a consistent fixed effects estimator of β for probit models. A more complete discussion of this point can be found in Hsaio [25].[7]

Table I reports results from estimating four specifications of Eq. (5). Each specification includes the full price of medical care, measures of health capital and related individual specific variables that may affect demand for medical care, and measures of air pollution levels. A likelihood ratio test statistic for joint significance of all explanatory variables appears beneath each specification. Several variables present in the theoretical model are excluded from the equations estimated. Total time available (T) during the sample period is the same for all respondents. Also, the money price of the composite good (X) and the time required to consume one unit of it are assumed to be identical across respondents as well. Thus, q_X is excluded from the analysis because it varies only with the wage rate (w), a variable included in q_M. Additionally, through construction of the full income budget constraint (Eq. (3)), the total income variable was eliminated. While the model could be reformulated to include nonlabor income as a component of

[7]Fixed effects logit models are consistently estimable. A potential disadvantage of the random effects model is the possible correlation of X_{it} and u_i. A correlation would arise naturally if X_{it} included choice variables and the individual knew his own u_i: utility maximization would make the choice of X_{it} dependent on u_i. Such a correlation should not affect the demand equations estimated below, since the equations are reduced forms with only exogenous explanatory variables.

TABLE I
Estimation Results and Descriptive Statistics[a]

Explanatory Variable[a]	Probit parameter estimates of the probability of seeking medical care[b] (1)	(2)	(3)	(4)	Sample means and standard deviation (5)
CONSTANT	−2.786	−2.611	−2.605	−4.026	
	(−3.042)	(−3.740)	(−3.717)	(−4.026)	
FPMED	−0.4615E-02	−0.4510E-02	−0.451E-02	−0.4606E-02	35.33
(dollars/visit)	(−1.480)	(−1.603)	(−1.602)	(−1.596)	35.64
DOCREG	0.3267	0.2905	0.2908	0.3089	0.8297
	(1.206)	(1.477)	(1.480)	(1.548)	0.3761
CHRONIC	−0.3228	−0.3106	−0.3410	−0.2995	0.2166
	(−1.382)	(−1.737)	(−0.994)	(−1.667)	0.4121
SHRTWHZ	−0.8606E-01	−0.7920E-01	−0.7929E-01	−0.8932E-01	0.3944
	(−0.4642)	(−0.5460)	(−0.547)	(−0.6130)	0.4890
FLEMCO	0.1652	0.1865	0.1858	0.1533	0.2532
	(0.8226)	(1.226)	(1.220)	(0.9910)	0.4351
HAYFEV	0.3568	0.3390	0.3388	0.3481	0.2144
	(1.766)	(2.152)	(2.150)	(2.188)	0.4107
BURB	0.3263	0.2992	0.2990	0.2995	0.2392
	(1.167)	(1.202)	(1.202)	(1.191)	0.4268
BLUE	−0.1776	−0.1694	−0.1683	−0.1743	0.3017
	(−0.8264)	(−1.014)	(−1.006)	(−1.033)	0.4593
EXPWORK	0.3660	0.3423	0.3414	0.3428	0.3955
	(2.193)	(2.420)	(2.409)	(2.404)	0.4892
EDGRADE	0.5107E-01	0.4605E-01	0.4595E-01	0.5021E-01	14.72
(years)	(1.597)	(1.666)	(1.663)	(1.789)	2.584
AGE	−1.1592E-02	−0.1446E-02	−0.1442E-02	−0.1541E-02	47.57
(years)	(−0.1392)	(−0.1650)	(−0.164)	(−0.1720)	7.686
CHRONIC	—	—	0.2855	—	1.988
×O3			(0.104)		4.642
O3	0.4960E-01	0.4755E-01	0.4708E-01	0.2584E-01	9.350
(pphm)	(3.852)	(3.643)	(3.367)	(1.396)	6.021
CO	−0.3482E-01	−0.2818E-01	−0.2810E-01	−0.2216E-01	3.429
(ppm)	(−0.6043)	(−0.5170)	(−0.516)	(−0.3930)	2.682
SO2	0.3406E-01	0.2136E-01	0.2119E-01	0.6638E-02	0.8317
(pphm)	(0.1300)	(0.1150)	(0.114)	(0.3500E-01)	0.4971
NO2	−0.1918E-01	−0.2000E-01	−0.1996E-01	−0.1296E-01	8.469
(pphm)	(−0.6760)	(−0.7460)	(−0.745)	(−0.4740)	3.679
TEMP	—	—	—	0.1702E-01	75.97
(°F)				(2.109)	15.47
HUMID	—	—	—	0.3926E-02	42.64
(percentage)				(0.7000)	15.88
σ_u^2	0.2618	—	—	—	
	(1.044)				
$-2 \ln LR$[c]	40.53	40.08	40.09	47.481	

[a]Units of measurement are given beneath variable names (for all variables that are not 0–1 dummies).
[b]Asymptotic t statistics in parentheses. The symbol "E-0n" refers to 10 raised to the $-n$th power.
[c]This test statistic is -2 times the log of the likelihood ratio, where the restricted likelihood is calculated by estimating the model with all slope coefficients zero.

full income, nonlabor income was not measured on the survey. Finally, measures of race, gender, and marital status were not included because nearly 90% of the respondents are married white males.

Specification (2) in the table is identical to specification (1) with the restriction that σ_u^2, the variance of the individual-specific error component, is zero. Taken together, these specifications allow a likelihood ratio test of the null hypothesis that there is no individual-specific variation in the probability of seeking medical care, after effects of included explanatory variables have been controlled for. The p value for the test is 0.56, indicating that the hypothesis cannot be rejected even at the relatively high significance levels often recommended for pretests.[8] Individual differences in unmeasured variables that determine the probability of seeking medical care (such as $\overline{M}$, the stock of health related information acquired from past medical contacts), therefore, appear to be small. As a consequence, welfare effects of air pollution changes are measured with the medical demand equation estimated by dropping the error components specification and using ordinary probit.[9]

In specifications (2), (3), and (4), the coefficients of FPMED and DOCREG are respectively negative and positive and are both significant at lower than 10% in a one-tail test, indicating that those respondents with a regular doctor and lower time and money costs are more likely to seek medical care. In unreported specifications, the three components of FPMED (PMED, TMED, WAGE) were entered individually. Coefficients of each were negative and those of TMED and WAGE were significantly different from zero.

The puzzling negative coefficient of CHRONIC apparently occurs because the respondents in this category sought medical care less frequently during the sampling period than other respondents, despite reporting higher typical and recent annual frequencies of doctor visits. Yet, the presence of physician diagnosed chronic lung disease appears to have no effect on the relationship between ozone pollution and medical care demand. In specification (3), coefficients of the dummy variable CHRONIC and an interaction term measuring the product of CHRONIC and O3 are individually and jointly insignificant at conventional levels, while remaining parameter estimates are essentially unchanged by the inclusion of the interaction term.

A broader examination of the effect of chronic lung disease on medical care demand was undertaken by reestimating Eq. (2) with a full set of interaction terms in addition to the dummy variable CHRONIC allowed for. The null hypothesis that the constant term and coefficients of all explanatory variables jointly are identical between individuals with and without physician diagnosed lung disease,

[8]As a precaution against premature simplification of the model, the test was repeated using a number of alternate specifications of explanatory variables, including those from specifications (3) and (4) of Table I, which includes temperature and humidity variables. The p value of the test is insensitive to specification changes, except those that restrict the individual-specific, time invariant variables in the equations. If most or all of the individual-specific variables except the price of medical care (DOCREG through AGE in Table I) are excluded, the null hypothesis can be rejected at conventional significance levels. Thus, it appears that the explanatory variables used in the equations control adequately for the individual-specific variation which the permanent error component is supposed to capture.

[9]Since u_i includes the effects of all omitted individual-specific, time invariant variables, failure to reject $\sigma_u = 0$ does not necessarily imply that current medical care is an essential input to the health production function.

tested using the likelihood ratio test procedure, cannot be rejected at the 10% significance level ($p = 0.15$).

Performance of remaining health status variables is mixed. Coefficients of dummy variables SHRTWHZ and FLEMCO are insignificant. In fact, if a full set of health dummies measuring compromised respiratory function (ASTHMA, BRONCH, SHRTWHZ, FLEMCO) is included in the equation, the null hypothesis that their coefficients are jointly zero cannot be rejected at significance levels between 35 and 50%. In contrast, the coefficient of HAYFEV, which would seem to be a much less serious ailment, is positive and significant.

Equations (2) through (4) also indicate that residents of Burbank, individuals who are exposed to substances at work that affect their breathing, the more highly educated, and those in white collar occupations are more likely to seek medical care. Age appears to have little effect on the decision to seek medical care, a result which is not surprising because the very young and very old are not represented in the sample.

Turning to the pollution variables, the coefficient of O3 is positive in all specifications and is significant at 1% in all specifications except (4), in which it is significant at 10% in one-tail tests. All other pollution measures have t-statistics less than 1 in absolute value, but the four pollution measures are jointly significant at 1% in all equations except (4). Thus, in Southern California, elevated ozone levels, as contrasted with elevated levels of other pollutants, appear to cause people to seek medical attention. This result is plausible because elevated ozone levels can cause immediate symptoms such as chest pain, throat irritation, sinus pain, and headache, although collinearity between the pollution variables is a possible concern.[10] The smaller coefficient of O3 and higher associated standard error in specification (4) as compared with those in specification (2) results from including the climate variables TEMP and HUMID. The Pearson correlation between TEMP and O3 is 0.428, an association that is expected because ozone is a secondary pollutant formed by the interaction of other pollutants in sunlight. As a consequence, ozone concentrations tend to be high when temperature is high, although it is less clear why higher temperatures rather than higher ozone levels would induce doctor visits when daily high temperatures average about 76°F. Nevertheless, the estimated effect of ozone on the probability of seeking medical care still is positive and significant at the 10% level with TEMP included in the equation. In Eq. (4), the coefficient of TEMP also is positive and significant at the 5% level using a one-tail test, while the coefficient of HUMID has a t ratio of less than unity.

To use Table I estimates to compute willingness to pay for improved air quality, define $y = \bar{x}\hat{\beta}$ as the inner product of explanatory variables and estimated coefficients, with each explanatory variable except ozone set equal to its sample mean, and let $F(\cdot)$ denote the standard normal cumulative distribution function. $F(y)$ is the estimated probability of obtaining medical care and is interpreted as the Marshallian demand function evaluated at $\bar{x}$. Let y^0 represent y evaluated with ozone set equal to a lower bound value of $O3^0$ (e.g., $O3^0 = 12$ pphm, the current federal standard), and let y^i represent a value of y where the peak ozone reading of $O3^i$ exceeds the lower bound. Following Small and Rosen [20], the change in

[10] Pearson correlations between ozone and other pollutants are 0.04 for CO, 0.26 for SO2, and 0.50 for NO2; the largest correlation between any pair of pollutants is 0.61 between CO and SO2.

TABLE II
Illustrative per Person per Year Medical Expense and Willingness to Pay Values for Ozone Control

Maximum peak daily ozone level (pphm)	City	Eq. (2) Consumers' surplus ($)	Eq. (2) Medical expense ($)	Eq. (4) Consumers' surplus ($)	Eq. (4) Medical expense ($)
12	Burbank	115	58	95	25
	Glendora	209	110	171	46
9	Burbank	205	90	171	41
	Glendora	314	148	261	65

consumers' surplus area behind the medical demand curve associated with reducing peak ozone levels from $O3^i$ to $O3^0$ is given by the integral

$$\mathrm{CS}\,(O3^i, O3^0) = -(1/\lambda)\int_{y^0}^{y^i} F(y)\,dy, \tag{7}$$

where λ denotes the marginal utility of income, which is factored out of the integral because it is a constant equal to the full price of medical care in the Table I specifications of Eq. (5). Because there is no closed form solution for $F(\cdot)$, the integral is approximated numerically using a Gauss–Konrad quadrature rule. Equation (7) will approximate willingness to pay accurately if the ordinary demand curve lies close to the compensated demand curve (see Small and Rosen [20]).

The willingness to pay estimate in Eq. (7) can be compared to the expected change in out-of-pocket medical expenditures by estimating the associated change in demand and multiplying by the money price: $\mathrm{PMED}[F(y^i) - F(y^0)]$. The change in medical expenses is one component of damages (another is value of lost work time) obtained when the cost of illness method is applied. Unlike willingness to pay to reduce pollution, the cost of illness is not a theoretically correct measure of benefits but nonetheless is widely used. An advantage of applying the valuation approach outlined above to medical care is that the theoretically preferable measure in Eq. (7) can be compared to one component of illness costs.

Willingness to pay and medical expense estimates are nonlinear functions of the upper and lower bound ozone values. For example, daily CS(30, 12) = \$4.06, while daily CS(12, 9) = \$0.64. Illustrative calculations were made using two lower bound ozone values: the current standard of 12 pphm and a more stringent goal of 9 pphm. Representative upper bound ozone values were chosen at 3-pphm intervals from a maximum of 30 pphm to a minimum of 3 pphm above the lower bound. Daily willingness to pay and medical expense changes then were calculated for each pair of upper and lower bound ozone readings and aggregated to annual values. To aggregate, each daily estimate was multiplied by the number of days in 1985 on which the peak ozone reading fell in the 3-pphm interval less than or equal to the relevant upper bound.[11]

Results of this illustrative valuation procedure are presented in Table II, while the 1985 frequency distributions of daily maximum one-hour ozone readings for

[11] One exception is that all ozone readings over 30 pphm were assigned the value 30.

TABLE III
Frequency Distribution of Daily Peak Ozone Levels: Glendora and Burbank, 1985

Peak ozone level (pphm)	Number of days	
	Burbank	Glendora
< 3	85	86
4–6	80	69
7–9	59	44
10–12	54	49
13–15	40	33
16–18	27	27
19–21	8	24
22–24	5	18
25–27	6	8
> 28	1	7

Burbank and Glendora used in the calculations are presented in Table III. Willingness to pay and medical expense estimates presented are measured in dollars per person per year (rounded to the nearest dollar) for an environment in which daily peak ozone levels never rise above 12 and 9 pphm on any day of the year. Separate calculations are presented for Burbank and Glendora and for specifications 2 and 4 reported in Table I (results from specification 3 are similar to those from 2).[12]

Both types of estimates are lower for Burbank and when based on Eq. (4). Table III shows that in 1985, Burbank had 30 fewer days than Glendora in which peak hourly ozone levels exceeded 12 pphm and 25 fewer days than Glendora in which the same air pollution measure exceeded 9 pphm. Also, Eq. (4) yields lower estimates than Eq. (2) because of its smaller coefficient of O3. Using the consumers' surplus figures, rough annual benefit estimates of meeting the current federal ozone standard on each day in 1985 range from \$171 to \$209 in Glendora and \$95 to \$115 in Burbank. If the federal standard were reduced to 9 pphm, corresponding estimates would rise by more than 50% in Glendora and about 90% in Burbank. Estimates of medical expenses are lower than those for consumers' surplus by factors of between 2 and 4 in all situations considered.

5. IMPLICATIONS AND CONCLUSIONS

Willingness to pay estimates presented are likely to be lower bounds on true values in view of problems in ensuring that the essential input condition is satisfied. Both willingness to pay and medical expense estimates would increase slightly if income taxes or other distortions caused the wage rate to exceed the marginal value of time used to calculate the time price of medical care. Willingness

[12]Differences in annual value estimates between Burbank and Glendora are caused solely by differences in the ozone frequency distributions. Separate daily calculations could be made for the two communities since BURB was included as an explanatory variable; however, this refinement was not pursued because the coefficient of the dummy variable BURB was not significantly different from zero at conventional levels.

to pay and the value of time move in opposite directions because higher values of time reduce ozone induced shifts in medical demand for all equations estimated (i.e., $\partial^2 F(y)/\partial(\mathrm{WAGE})\partial(\mathrm{O3}) < 0$). Empirically, each 10% reduction in the marginal value of time leads to a 1 to 3% increase in consumers' surplus estimates.[13]

Although the approach taken here may understate willingness to pay for reduced tropospheric ozone levels, consumers' surplus estimates presented in Table II are larger than most related estimates obtained in previous studies of oxidant pollution and health. For example, the estimates above are much larger than those obtained in damage function analyses. Seskin [26] concluded that a 50% reduction in maximum one-hour oxidant levels in Washington, D.C., in 1973–1974 would reduce medical expenditures by about $0.04 per person per year. Portney and Mullahy [27], on the other hand, found that a 10% nationwide reduction in average daily maximum one-hour ozone readings in 1979 would have resulted in between 0.25 million and 22 million fewer respiratory related restricted activity days (RRADs) among adult residents of U.S. urban areas. Valuing an RRAD at $20 per day results in per person per year benefit figures ranging from $0.04 to $4.00.

Differences between consumers' surplus and damage function estimates cannot be reconciled completely; however, three possible explanatory factors are worth citing. First, Seskin's analysis focuses only on medical expenses, ignores disutility effects and value of lost work time, and therefore underestimates willingness to pay. Second, the panel used here includes a disproportionately large number of respiratory impaired respondents. For example, approximately 22% of the individuals in the panel report physician diagnosed chronic lung function impairment, while only 17% of the Portney and Mullahy sample report chronic impairments of any type. Yet, the estimated difference in the demand response to ozone changes between individuals with and without chronic lung disease is negligible both in magnitude and in statistical significance (see Eq. (3) of Table I). Differences in surplus exist only when differences in demand exist, but significant demand differences were found only in the constant term and only by imposing equality of coefficients of explanatory variables. Since the constant shift presumably reflects the previously mentioned disparity between relative numbers of typical and in-sample doctor visits for respondents with and without chronic lung disease, the basis for making separate welfare calculations for the two groups appears weak. Nevertheless, if separate calculations are made, values for individuals with chronic lung disease are lower than those reported in Table II, while values for those without chronic lung disease are higher. Evidently, the respiratory impaired respondents are not the source of the relatively high willingness to pay estimates reported in Table II.[14]

[13]Browning and Johnson [32] estimate that the average U.S. worker faces an overall marginal tax rate of 0.43. Assuming a marginal value of time of $0.6 \times \mathrm{WAGE}$ and using Eq. (2) estimates from Table I, willingness to pay estimates for a 12 pphm ozone environment rise to $125 in Burbank and $227 in Glendora.

[14]Using Eq. 2 estimated from Table I, consumers' surplus estimates computed using sample means of explanatory variables for individuals with CHRONIC = 1 are 52 to 53% of estimates in Table II for both cities and both lower bound ozone values. Estimates for those with CHRONIC = 0 are 18 to 19% larger than Table II estimates.

A third and possibly the most important explanation for the discrepancy between willingness to pay and damage function estimates is that observed ozone levels were quite low in both the Seskin and Portney and Mullahy studies. In Seskin's study, average maximum one-hour ozone concentrations ranged from 3.2 to 4.8 pphm, depending on which monitoring station and year are considered. Also, the national primary and secondary oxidant standard of 8 pphm in force between 1971 and 1978 was violated only 48 times in 1973. In the Portney and Mullahy study, the sample mean of the average daily maximum one-hour ozone reading was 4.2 pphm. Corresponding ambient ozone readings from Glendora and Burbank, as shown in Table I, were 9.3 pphm, which is two to three times larger than the figures just listed, and as shown in Table III, 13% of the readings in Burbank and 23% of the readings in Glendora exceeded 15 pphm. Moreover, epidemiological and medical studies generally do not find measurable health effects of oxidant pollution until levels rise to the 8- to 10-pphm range (for a review of this evidence, see Gerking *et al.* [28]).

Contingent valuation and household production studies have obtained larger estimates of willingness to pay for reduced ozone levels than those found by Seskin and Portney and Mullahy. Schulze *et al.* [29], for example, asked survey respondents in the Los Angeles area to recall a highly publicized ozone episode and, through a series of contingent valuation questions, found that willingness to pay averaged about $7.75 per person per day to reduce peak one-hour ozone concentrations from 20 to 12 pphm. This outcome translates into conservative annual per person willingness to pay estimates of about $403 per resident of Glendora and $132 per resident of Burbank obtained by multiplying $7.75 by the number of days that one-hour peak ozone levels exceeded 20 pphm (see Table III). These values, which take no account of days when peak ozone levels are between 12 and 20 pphm, are larger than the consumers' surplus estimates reported in Table II. Also, in their study of the role of medical care in home producing health, Gerking and Stanley [14] calculated that residents of St. Louis (a city with lower tropospheric ozone pollution levels than Los Angeles) were willing to pay about $24 per year each for a 30% reduction in overall ozone exposures. In another household production study that used the same Glendora/Burbank data set and extended the analytical framework of Joyce *et al.* [17], Dickie and Gerking [18] found that adults with normal respiratory function are willing to pay about $75 annually to avoid days on which peak ozone levels exceed 12 pphm.

This broad range of willingness to pay estimates poses an awkward situation for policy makers because more stringent measures to control tropospheric ozone pollution currently are under consideration both at the federal level and in California. In the summer of 1989, President Bush presented a plan to Congress that included incentives for automakers to manufacture engines that operate on alternative fuels such as methanol as well as controls on gasoline pump nozzles to prevent fumes from escaping into the atmosphere. In California, where ozone pollution is a relatively more serious problem than in other parts of the United States, the 1989 Air Quality Management Plan for the South Coast Air Basin contains more comprehensive recommendations [30]. Selected measures include controlling emissions of reactive organic gases from: (1) solvents and coatings used to refinish and degrease wood, automobiles, marine vessels, and aerospace equipment, (2) transporting and dispensing gasoline products, (3) aerosol antiperspirants, and (4) commercial charbroiling by fast food and full service restaurants. To the extent that the larger willingness to pay estimates such as those found in this study are valid, more aggressive measures to control tropospheric ozone pollution would be warranted, particularly in California.

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