Measuring the Cost of Time in Recreation Demand Analysis: An Application to Sportfishing

Kenneth E. McConnell and Ivar Strand

We reckon hours and minutes to be dollars and cents.
—T. C. Haliburton, *The Clockmaster*

Since the work of Cesario and Knetsch, economists have recognized that the opportunity cost of time plays an important role in determining the demand for outdoor recreation. The opportunities one has for spare time are more significant for consumption of time-intensive outdoor recreation activities than for other commodities, especially nondurables. Bishop and Heberlein illustrate "the overwhelming importance of time costs to final [recreational] values. . . . Total consumer surplus is nearly four times as large . . . [when] time costs are added at half the income rate . . . [as when] time costs were set at zero" (p. 21).

Despite the recognition, economists have neither successfully integrated the costs of time with the methods of recreational demand analysis nor reached a consensus on how it should be measured. Brown, Charbonneau, and Hay state, "Finally, the apparently crucial importance of how opportunity cost of time is handled needs further work. While we are convinced it is an appropriate concept, . . . exactly how it should be included and measured . . . remains to be determined" (p. 24). Several approaches have been taken to include it in the travel cost method. One approach (Brown and Nawas, Gum and Martin) suggests that time in transit be considered as a separate independent variable. Another approach (Bishop and Heberlein; Brown, Charbonneau, Hay; Nicols, Bowes, Dwyer; Cesario and Knetsch) measures the cost of time and adds it to other costs. Several approaches have been suggested to measure time costs. One approach is simply to choose an hourly wage, e.g., $2.00 per hour, or perhaps the minimum wage rate. A more flexible but still ad hoc approach is to use some proportion of the individual's wage rate as the opportunity cost of time (Nichols, Bowes, Dwyer). The proportion is usually taken from independent studies and used to value the travel time. This approach is better than using a constant opportunity cost of time because it allows variation across individuals. It suffers because the choice of the percentage of the wage rate is arbitrary, independent of the sampled population. Cesario has discussed the consequences of ignoring time costs and the differences in values arising from alternative measurement approaches.

In this paper, we argue that the opportunity cost of time is some proportion of the individual's market wage rate or income per hour and that this proportion can be determined from sample data. This method permits the proportion to vary from one study to another, rather than imposing either an arbitrary estimate or one from a sample different from the study's sample.[1]

A Simple Model

The recreationist presumably behaves as if to maximize utility subject to time and budget constraints by choosing trips, denoted r. The original travel cost method (Clawson) used trips per capita (z) as the dependent variable. In this paper, we have chosen to use trips per user (r). But $z = \Pi r$, where Π is the participation rate (proportion of population who participate at least once). Various studies (e.g., Deyak and Smith) have shown that decisions to participate are different from decisions about how frequently to participate. As Brown and Nawas point out, there is loss of information in aggregation. Hence it is more efficient to use r as a dependent variable. However, the method we discuss will work for z or r as the dependent variable.

Let utility be $U(x,r)$, where r is recreation trips and x is a bundle of all other goods. If we introduce a proportionate income tax rate of t, the budget constraint is

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Kenneth E. McConnell and Ivar Strand are, respectively, associate professor and assistant professor in the Department of Agricultural and Resource Economics, University of Maryland.

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[1] The method as described is similar in spirit to a method described in Commons. This paper, brought to our attention by a reviewer of a version of this paper, describes a method of choosing the proportion for a log-linear demand function by a search method.

$$(1) \qquad [F(w) + E](1 - t) = px + cr,$$

where w is the amount of time worked, $F(w)$ is income earned from w units of work, E is fixed income, t is the income tax rate, p is the price of the composite bundle, and c is out-of-pocket costs per recreational trip. Before-tax income is $F(w) + E$. It is the most frequent measure available from surveys. Suppose the time constraint is given by $T = ar + w$, when T is total time available and a is the amount of travel time per recreational trip.[2] The problem is to maximize

$$(2) \quad U(x,r) - \lambda\{px + cr - (1 - t)[F(T - ar) + E]\}.$$

The first-order condition for r is

$$(3) \qquad \partial U/\partial r = \lambda[c + a(1 - t)F'(w)].$$

Assuming that p does not vary across individuals, we get the demand function for recreation:

$$(4) \qquad r = f[c + a(1 - t)F'(w)].$$

Income is given by $F(w) + E$, while the marginal opportunity cost of time is $(1 - t)F'$. Define average income by $v = [F(w) + E]/w$. Sufficient conditions for the cost of time [measured by $(1 - t)F'$] to equal v are (a) The tax rate, t, is zero; (b) marginal earnings are constant: $F'(w) = F(w)/w$; and (c) nonwork income, E, is zero.

From these, it appears likely that the opportunity cost of time is less than average income. If the income figure is family income where other family members earn income and $v \equiv$ family income/w, the individual's opportunity cost of time will be overstated. The opportunity cost of time will be understated if an individual gets utility from work or if working today is a form of investment which provides higher income in the future.

Suppose the opportunity cost of time is some constant (k) times the average income. Then the demand function is

$$(5) \qquad r_i = f(c_i + ka_iv_i),$$

where $0 < k < 1$ is usually an arbitrarily chosen number and i is an observation index. Instead of choosing k arbitrarily, we let the sample determine k. With a linear form, we have

$$(6) \qquad r_i = \beta_0 + \beta_1(c_i + ka_iv_i) + \beta_3Z_i + e_i,$$

where Z_i is a vector of exogenous variables including a wealth or income proxy and e_i is an error term with the classical specification. We can rewrite (6) as

$$(7) \qquad r_i = \beta_0 + \beta_1c_i + \beta_2a_iv_i + \beta_3Z_i + e_i.$$

[2] We assume that a is travel time per trip. This approach implies that the opportunity cost of time spent on site is zero. While this is standard practice (Brown and Nawas; Shulstad and Stoevener), it is an unresolved but important issue (McConnell). We do not attempt to deal with the issue in this paper.

The estimate of k is $\hat{k} = \hat{\beta}_2/\hat{\beta}_1$, where $\hat{\beta}_i$ are the ordinary least squares (OLS) estimates of the parameters of (7). In the following section we show how this method works on a sample of sportfishermen.

An Application to Sportfishing

To test the approach suggested, we use sample data from a 1978 survey of sportfishermen in the Chesapeake Bay region. The complete specification of the equation is

$$(8) \quad r_i = \beta_0 + \beta_1c_i + \beta_2a_iv_i + \beta_3s_i + \beta_4m_i + e_i,$$

when r is the annual sportfishing trips per angler, c is per trip expenses per person, a is the round trip travel time (computed as round trip distance/45 miles per hour), v is average hourly income (annual family income/2080), s is a site variable equaling 1 for residents of Ocean City, Maryland, and 0 otherwise, and m is the length of the angler's boat.

The expected signs and relationships are $\beta_1 < \beta_2 < 0$, $\beta_3 > 0$, $\beta_4 > 0$. The first two inequalities relate to the negative effect of costs, both trip expenses and travel time, on the trips taken per year. Also, $\beta_1 < \beta_2$ implies that the opportunity cost of travel time is less than average income. The site variable (s) attempts to capture variation due to different characteristics of the sites. Since Ocean City, Maryland, was our only resort area, it was given a value of 1 and the other sites given 0. Boat length (m) represents a previous commitment to sportfishing or a wealth proxy. In either case, it should act to increase annual participation.

Fitting equation (8) on the Maryland-Virginia survey gives us

$$(9) \; r = 9.77 - \underset{(3.89)^*}{.0206c} - \underset{(2.00)}{.0126av} + \underset{(2.50)}{.019s} + \underset{(5.06)}{.157m},$$

where $N = 415$, $\bar{R}^2 = .10$, $F(4,411) = 12.8$, and asterisk indicates t-statistics under the null hypothesis of no association. For this equation we have used a subset of observations from the sample.[3] The estimated coefficients agree in sign and magnitude with our prior beliefs. The equation fits reasonably well for cross-sectional observations.

[3] The subset of the sample included anglers who made twenty or fewer trips per season. To test whether the groups were different, a Chow test was used. The test statistic [$F(234,412) = 27.3$] permitted rejection at the 99% confidence level of the null hypothesis that the coefficients of the equation (9) were the same for anglers with twenty or fewer trips and anglers with more than twenty trips. We report results only for the twenty or fewer group. The hourly income variable was based on seven annual income categories ($0–$4,999; $5,000–$9,999; $10,000–$14,999; $15,000–$19,999; $20,000–$29,999; $30,000–$49,000; $50,000 and above) with the average of the category range being assigned to respondents in the category. No respondents from the lowest range were used because respondents not wishing to reveal their income often responded by indicating the lowest income class. This exclusion limits the range of v but appeared more appropriate than introducing considerable error and biased data by inclusion. For a detailed description of the survey, see Strand and Yang.

Using equation (9), we can infer that a representation angler values time at about 60% of his hourly income:

(10) $\hat{k} = \hat{\beta}_2/\hat{\beta}_1 = -.0126/-.0206 = .612.$

We expect that k will vary among regions and sites and that this value is applicable only to our sample. However, by estimating it directly from observations on individual behavior we have eliminated the need for ad hoc and arbitrary valuation of the opportunity costs of time.

Properties of $\hat{k}$

As we have observed, variations in k cause considerable variations in estimates of consumers' surplus. Our value of $\hat{k}$ is not the true value but rather the ratio of two random variables; hence, it is a random variable itself. The reliability of the estimate of consumers' surplus depends on the random properties of $\hat{k}$.

We can ascertain something of the underlying probability distribution of $\hat{k}$ from what we know of $\hat{\beta}_2$ and $\hat{\beta}_1$. Under classical assumptions, the distribution of these coefficients is jointly normal. The distribution of the ratio of two $N(0,1)$ variables is a standard form Cauchy (Johnson and Kotz, chap. 16). However, if the variables forming the ratio are jointly dependent, as are $\hat{\beta}_1$ and $\hat{\beta}_2$, then the underlying distribution is more complex (Springer, chap. 4). In both cases, however, the distributions do not have finite moments. Since confidence intervals and significance tests rely on the existence of second moments, neither of the traditional tests is applicable. We can develop some understanding of the dispersion of $\hat{k}$ by Monte Carlo studies of the ratio of jointly normal variates. This procedure offers guidance about the distribution of $\hat{k}$.

Let the joint density function of $\hat{\beta}_1$ and $\hat{\beta}_2$ be given by $f(\hat{\beta}_1,\hat{\beta}_2)$. Then

(11) $$f(\hat{\beta}_1,\hat{\beta}_2) = f_1(\hat{\beta}_1)f_2(\hat{\beta}_2|\hat{\beta}_1),$$

where $f_1(\hat{\beta}_1)$ is the marginal density function of $\hat{\beta}_1$, and $f_2(\hat{\beta}_2|\hat{\beta}_1)$ is the conditional density function of $\hat{\beta}_2$ given $\hat{\beta}_1$. With these conditions, it can be shown that

(12) $$\hat{\beta}_1 \sim N(\beta_1,\sigma_1^2), \text{ and}$$

(13) $$\hat{\beta}_2|\hat{\beta}_1 \sim N[\beta_2 + \rho\sigma_2(\hat{\beta}_1 - \beta_1)\sigma_1^{-1},\sigma_2^2(1 - \rho^2)],$$

where ρ is the correlation coefficient of the bivariate normal. With conditions (12) and (13) we can construct two random variables which follow (11) by calculating

(14) $$\tilde{\beta}_1 = \beta_1 + \sigma_1\Theta_1,$$

(15) $$\tilde{\beta}_2 = \beta_2 + \sigma_2[\Theta_2(1 - \rho^2)^{\frac{1}{2}} + \Theta_1\rho], \text{ and}$$

(16) $$\tilde{k} = \tilde{\beta}_2/\tilde{\beta}_1,$$

where Θ_i are $N(0,1)$ and independent. We performed experiments by drawing sequential pairs of unit normal random variables, assuming that the true value of β_1, β_2, σ_1, σ_2 and ρ were as estimated in equation (9). The assumed values are $-.0206$, $-.0126$, .0067, .0050, and $-.3781$, respectively.

Several experiments with sample size varying from 50 to 1,000 were conducted (table 1). Each row gives the mean value of $\tilde{k}$, the bias $(\hat{k} - \tilde{k})$, $\hat{k}$ being the ratio of estimated coefficients, the proportion of estimates greater than zero, and the proportion of estimates in the unit interval. Based on all experiments, there is an estimated probability of .016 that the estimates of $\tilde{k}$ will be less than zero. Our experiments also show that 66.7% of the sample ratios fell in the unit interval.

Although these results do not have the theoretical support of formal confidence intervals, they are informative. Despite the possibility of substantial dispersion as $\tilde{\beta}_1$ approaches zero, the experiments show remarkable conformity with the distribution of estimates. Though we cannot say $\tilde{k}$ is significantly different from zero at the 98.4% level of confidence, it seems reasonable to reject the hypothesis that the ratio is less than or equal to zero.

The alternative to the Monte-Carlo approach is to assume that $\hat{k}$ is asymptotically normal with expected value $E\hat{\beta}_2/E\hat{\beta}_1$ and variance approximated by

(17) $$V(\hat{k}) = (\hat{\beta}_2/\hat{\beta}_1)^2[\hat{\sigma}_1^2/\hat{\beta}_1^2 + \hat{\sigma}_2^2/\hat{\beta}_2^2 - 2\,\text{cov}(\hat{\beta}_1,\hat{\beta}_2)/\hat{\beta}_1\hat{\beta}_2].$$

Using the values of their variables following equation (16), we compute $V(\hat{k}) = .142$. With these assumptions and numbers, we can construct the standard rejection region for the null hypothesis that $k = 0$. For a type 1 error of 10%, the critical region for rejection of the null hypothesis lies beyond .483. Thus, based on this approximation, we would reject the null hypothesis that $k = 0$ because the estimated value of k is .612.

The comparison of the assumption of normality with the Monte-Carlo results indicates the kinds of errors we make by assuming normality. Under the condition that $\hat{k}$ is $N(.612, .142)$, about 80% of observations so distributed will fall in the unit interval, compared with about 67% from the Monte-Carlo results. Thus, this assumption of normality with mean .61 and variance given by (17) leads to underestimating the type 1 error. This difference suggests care in the interpretation of results.

Table 1. Some Properties of $\hat{k}$ from Sampling Experiments

Sample Size	Mean Value	Bias of $\hat{k}$	Relative Frequency $\hat{k} \geq 0$	Relative Frequency $0 \leq \hat{k} \leq 1$
50	.765	−.154	.984	.672
500	.886	−.247	.983	.662
1,000	.763	−.152	.985	.667

Conclusions

This paper offers a method of estimating the opportunity cost of time in the demand for recreation. It can be used simultaneously with travel cost analysis, requiring only the interviewee's wage rate or income as additional data. It eliminates the need to rely on an exogenous estimate of the opportunity cost of time.

We have applied this technique to linear demand curves, and with linear functions, OLS provides direct estimates of the proportion. The general approach of letting the sample data choose the proportion is applicable to any functional form via the use of maximum likelihood techniques. An advantage of estimating k directly by maximum likelihood methods is that its asymptotic properties are well known.

The opportunity cost of time is determined by an exceedingly complex array of institutional, social, and economic relationships, and yet its value is crucial in the choice of the types and quantities of recreational experiences. Because of its complexity, one must be cautious in explaining it simply, as we have. In particular, while this method has promise, the measurements are not inconsistent with several competing hypotheses. For example, income per hour as time cost may reflect a negative income effect for sportfishing or the effect of income on the willingness to pay to avoid travel. In addition, this simple approach cannot explain why the opportunity cost of time is related to income for individuals working fixed hours.

Although this paper suggests a new direction, there are undoubtedly more advances to be made. For example, this method requires that the ratio of the opportunity cost of time to income per unit of time be constant for all sample observations. A significant improvement would be to let this ratio change as a function of leisure time or occupation.

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