View metadata, citation and similar papers at core.ac.uk
brought to you by CORE
provided by Research Papers in Economics

Valuing the recreational benefits from the creation of nature reserves in Irish forests.

Riccardo Scarpa[1], Susan M. Chilton[2], W. George Hutchinson[3] and Joseph Buongiorno[4]

([1]) Universita` degli Studi della Tuscia, Viterbo, Italy and Centre for Research on Environmental Appraisal and Management (CREAM), University of Newcastle Upon Tyne, United Kingdom.

([2]) CASPER, Department of Economics, University of Newcastle Upon Tyne, United Kingdom.

([3]) Department of Agricultural and Food Economics, Queens University Belfast, United Kingdom.

([4]) Department of Forest Ecology and Management, University of Wisconsin, Madison, United States of America.

Correspondence to be addressed to:

Riccardo Scarpa - Research Associate, Centre for Research in Environmental Appraisal & Management, Room 221, Department of Agricultural Economics and Food Marketing, University of Newcastle, Newcastle Upon Tyne, NE1 7RU, United Kingdom.
Tel. 0191 - 222 65 74 Fax. 0191 - 222 67 20 e-mail : scarpar@unitus.it

Abstract

Data from a large scale contingent valuation study are used to investigate the effects of forest attributes on willingness to pay for forest recreation in Ireland. In particular, the presence of a nature reserve in the forest is found to significantly increase the visitors' willingness to pay. A random utility model is used to estimate the welfare change associated with the creation of nature reserves in all the Irish forests currently without one. The yearly impact on visitors' economic welfare of new nature reserves approaches half a million pounds per annum, exclusive of non recreational values.

Keywords: Non-market Valuation, Contingent Valuation, Forest attributes analysis, Nature Reserves.

Introduction

Creating nature reserves (henceforth NRs) in public forests is one important way of preserving biological diversity and providing ecological goods to the public. Yet the economic benefits of the creation of NRs are not well known. Managers of public forests must often provide timber revenues as well as biodiversity protection and a natural setting for outdoor recreation. In much public woodland the managerial task is therefore that of providing both market and non-market goods. Although the creation of NRs in forests is sometimes in conflict with the use of woodland for outdoor recreation, a nature conservation site within the forest adds to most visitors' recreational experience. Some studies indicate that social benefits for non-market goods of forests are sizeable and may exceed those provided by traditional forest market products (i.e. Lockwood *et al.*, 1993).

The costs of creating a NR in a public forest may have an easy definition in terms of foregone timber revenue or shadow prices arising from the constraints imposed on the process of timber production. In contrast, defining and measuring the full social benefits produced by creating a NR is a challenge. The full value, economic or otherwise, of a NR is unlikely to fully represent the complexity, and often uniqueness, of functions supported by the ecosystem that the NR is designed to preserve. Differing ethical beliefs about the adequacy of monetary measures to represent ecological values add to the challenge. Yet, forest managers need to deal with these hard decisions and they are often required by policy makers to document the expected public benefits of conservation initiatives. Although total conservation values are very controversial, some aspects seem more amenable to economic analysis. This study explores the effect of creating NRs on the recreational value of woodlands.

NRs within public forests are areas of conservation landuse, mostly covering sites no greater than 10 to 20 hectares. They conform to two major types, unmanaged deciduous woodland including scrub woodland, and bog and wetland areas. Such conservation areas can support rare varieties of flora such as sphagnum and bryophyte communities while providing habitats for rare as well as common forms of wildlife (Mc Curdy, 1989). The conservation value of most forests can be enhanced by cooperation between foresters and ecologists. A conservation plan can be drawn up highlighting wildlife habitats, identifying fragile or unusual flora and integrating conservation practices in the forest management plan. If appropriate, a NR area can be designated to protect and conserve existing valuable ecosystems or to encourage their re-establishment. Within a NR area created in this way conservation management will take precedence over commercial forestry (Forest Service. 1991).

The use of non-market valuation methods is well established in the estimation of various economic values associated with environmental functions of forests and forest ecosystem. In particular, several authors have attempted to quantify benefits from individual forest attributes with non-timber valuation methods (Englin and Mendelsohn, 1991; Mattson and Li, 1995; Boxall *et al.* 1996). Of the multitude of functions performed by NRs we focus exclusively on forest recreation. This is, we believe, one of the first attempts to estimate the effect of the creation of NRs on willingness to pay (henceforth *WTP*) for recreational visits to public woodland. To do so, following McFadden (1974) and Hanemann (1984, 1989), we develop a probabilistic model to link discrete choice contingent valuation (henceforth CV) responses to forest attributes and the socio-economic characteristics of visitors. Using a random utility difference interpretation of the observed responses we define the distribution of *WTP* for visiting forests. A particular strength of the forest-attribute random utility function developed in this study is that it is based on broad CV data obtained from 26 Irish forests, involving nearly ten thousand visitors.

We find that *WTP* for visiting forests depends significantly on forest attributes, the presence of NRs being a prominent one. We use the empirical model to illustrate how to derive the distributions of *WTP* for the visitors to each forest, which are equivalent to a forests' access charge schedules. We then estimate forest-specific welfare changes that would result from establishing NRs in those forests currently without one. The estimated benefits of new NRs, from recreation alone, are substantial.

Extending the standard model of *WTP* estimation from CV models

CV has become one of the most widely used tools to estimate non-market public benefits from changes in environmental quality. Since its inception (Bishop and Heberlein, 1979) the referendum format, asking a specific amount for the *WTP*, has gradually grown in popularity. Because the respondents are only required to provide a Yes-No answer to a given *WTP* amount, this format relies on a smaller cognitive effort than the earlier open-ended format where the respondent was required to state a maximum WTP value. For this and other properties (Randall and Hoehn, 1987) the referendum format is now considered the best approach to elicit value responses in CV studies and its use was advised by the Blue Ribbon Panel for studies aiming at compensatory litigation for environmental damage assessment (NOAA, 1993). Thus, this was the format employed in the CV surveys conducted in the 26 recreational forests of this Irish study. A strong limitation of this format, however, is its relative sample inefficiency. Many observations are needed to get precise benefit estimates, especially when conditional estimation is the objective. For this reason, some researchers have investigated the performance of alternative approaches to estimate conditional probabilities, and the associated utility functions, to measure welfare changes. For example, Boxall *et al.* (1996), mutuating from marketing research, estimate visitors' utility function from choice experiments and compare the implied welfare estimates to those from a CV study. Choice experiments require the respondent to compare non-dominated alternative arrangements of forest attributes in a series of repeated choices. These observed pairwise choices are then used to estimate the parameters of a random utility difference function, ignoring the dependency created by repeated choices. The approach used in the present study, instead, simply relates the probability of positive response to a given bid amount to the levels of forest attributes experienced during the course of the forest visit.

The typical referendum CV design splits the random sample in K sub-samples each of which is probed by assigning a given bid amount t_k. The probability of a yes response at each bid amount is then estimated on the basis of the frequency of observed "Yes" responses at each bid amount.

In this study the object of interest is a structural relationship for $WTP|t$. The objective of estimation includes a parameter vector θ and the distribution of $WTP|t,\mathbf{x},\theta$, where $\mathbf{x}$ is a vector of relevant covariates. The most common way to estimate parametrically the measures of welfare change from dichotomous responses is to fit a linear index to a parametric cumulative distribution function (henceforth cdf). This linear index consists of the bid amount (or a transformation thereof) and a constant. To obtain conditional probabilities of Yes-No responses other socio-economic covariates can also be included in the model. The coefficient of the linear index can be linked to economic theory and interpreted as a random utility-difference function (Hanemann,1984,1989) or, via a simple reparametrization, as a valuation function (Cameron and James, 1987; Cameron 1988; McConnell, 1990). In either case θ is most commonly estimated by maximizing the sample log-likelihood function and the statistical properties of the model are identical. An often-used specification that seems to fit most data sets well and has appealing computational (concave sample likelihood function) and theoretical (non negativity of *WTP*) qualities is the probit or logit specification of a linear index with a natural log transformation of the bid amount. Other, less frequently employed distributions are the Weibull and Gamma, which are asymmetric and limited to the non-negative orthant. Often, though, a simple natural logarithm transformation of the bid amount, provides both a good fit of the observed responses and the often required non-negativity and asymmetry of the *WTP* distribution. The natural log transform was the original specification employed by Bishop and Heberlein in their seminal paper in 1979. After 20 years of CV applications it probably remains one of the most frequently adopted specifications and fits most data sets well (Sellar *et al.,* 1986; Downing and Ozuna 1996; Langford *et al.,* 1998; Ready and Hu, 1995 amongst others).

With a random utility theory interpretation of this specification (Hanemann and Kanninen 1996) the probability of observing a Yes response can be linked to the respondent's *WTP* for the proposed change, as follows. The visitor regards the enjoyment of the outdoor experience in the forest as a deterministic event, while for the analyst the determinants of utility derived from the visit are assumed to be observable only in part. These observable components constitute the deterministic part of the model. What determines the remainder of the utility level is unobservable to the analyst who assumes it to be stochastically distributed according to some given properties, which are typically summarised into a specific functional form.

In the particular case of forest recreation the deterministic component can be partitioned into two vectors of variables **q** and **z**. The vector **q** collects forest-specific attributes affecting the outdoor recreational experience and determining the site's recreational quality. The vector **z** collects socio-economic variables which characterize the visitor's idiosyncratic profile.

Let's define the observable component of the utility from the visit as $u(m\text{-}t,\mathbf{q},\mathbf{z})$ where m is the visitor's income, and t is the proposed access charge, and assume that the unobservable component is ε^{1}, so that the total utility level is :

$$u(m-t,\mathbf{q},\mathbf{z}) = v(m-t,\mathbf{q},\mathbf{z}) + \varepsilon^{1} \quad (1.)$$

Similarly, for the utility level in the absence of the visit we have:

$$u(m,\mathbf{z}) = v(m,\mathbf{z}) + \varepsilon^{0} \quad (2.)$$

The individual would agree to pay the amount t as an admission charge only if:

$$v(m-t,\mathbf{q},\mathbf{z}) + \varepsilon^{1} \geq v(m,\mathbf{z}) + \varepsilon^{0} \quad (3.)$$

Rearranging the arguments and using Δ to indicate the differences we have:

$$v(m-t,\mathbf{q},\mathbf{z}) - v(m,\mathbf{z}) \geq \varepsilon^{0} - \varepsilon^{1} \rightarrow \Delta v = \Delta\varepsilon \quad (4.)$$

In terms of probability, using the definition of cumulative distribution function $F_{\eta}(\cdot)$ for the event $\Delta\varepsilon$, the probability of observing a "Yes" response at a given bid amount t is:

$$Pr(\text{Yes}|t,\mathbf{q},\mathbf{z})=Pr(\Delta\varepsilon \leq \Delta v)= Pr(\eta \leq \Delta v)\equiv F_{\eta}(\Delta v) \quad (5.)$$

$$\rightarrow Pr(\text{No}|t,\mathbf{q},\mathbf{z}) = 1 - F_{\eta}(\Delta v) \quad (6.)$$

This is the well known random utility difference interpretation popularized by Hanemann (1984) of the dichotomous choice response to CV elicitation questions. However, in our case we estimate the utility function from M forests, each with a different combination of recreational features **q**. The recreational experience in one forest is similar, but not identical to that in another because each has a peculiar combination of recreational attributes $\mathbf{q}_j$. For the classic single bound case a maximum likelihood estimate can then be obtained by maximizing the following modified sample log-likelihood over the space of the parameters in Δv:

$$\Sigma_j\Sigma_i I_{ji} ln[1-F_\eta(\Delta v(t,\mathbf{q}_{ji},\mathbf{z}_{ji}))] + \Sigma_j\Sigma_i(1-I_{ji}) ln[F_\eta(\Delta v(t,\mathbf{q}_{ji},\mathbf{z}_{ji}))],\ i= 1 \quad N,\ j = 1 \quad M. \quad (7.)$$

where I_{ji} =1 if visitor i expressed the *WTP* the amount t in forest j.

Improving estimate efficiency and model specification

To improve on the well-known problem of sample inefficiency from CV referendum data, the *WTP* question is sometimes reiterated at a lower or a higher bid amount, depending on the outcome of the first response (the so-called “follow-up”, Hanemann *et al.* 1991). This additional response is often assumed to be generated by the same underlying *WTP* distribution as the first one, allowing interval data estimation of the probability model. Yet, the second response is clearly not independent from the first one and this may justify the use of bivariate estimation, where the first and second response are treated as being generated by two correlated, but distinct *WTP* distributions (Cameron and Quiggin, 1994). However, estimation of interval data models in Monte Carlo experiments run on responses generated by bivariate processes have shown that this assumption causes only a small bias, while increasing efficiency significantly (Alberini, 1995). For this reason we assumed that both responses were generated by the same underlying *WTP* distribution and use interval data analysis.

When the respondent is not aware that there will be a follow-up question, this format allows estimation of a probability model on the basis of the first response alone. This allows the researcher to fall back on the single bound estimates if the data actually provide evidence of strategic behavior in the follow-up responses.

To improve estimation efficiency in the Ireland forest recreation study each initial response was followed by a question with a bid amount (t^h) higher than the first one (t) when the first response was "Yes", and lower (t^l) when it was "No".

We therefore have two responses from which to estimate the distribution of Δv and the associated *WTP*. Under the assumption that the first and second response have the same underlying distribution of *WTP,* the interval data probabilities of the four possible responses are:

$$Pr(\text{Yes,Yes}|t,\mathbf{q},\mathbf{z}) = 1\text{-}F_\eta(\Delta v(t^h,\mathbf{q},\mathbf{z})) \quad (8.)$$

$$Pr(\text{Yes,No}|t,\mathbf{q},\mathbf{z}) = F_\eta(\Delta v(t^h,\mathbf{q},\mathbf{z})) - F_\eta(\Delta v(t,\mathbf{q},\mathbf{z})) \quad (9.)$$

$$Pr(\text{No,No}|t,\mathbf{q},\mathbf{z}) = F_\eta(\Delta v(t^l,\mathbf{q},\mathbf{z})) \quad (10.)$$

$$Pr(\text{No,Yes}|t,\mathbf{q},\mathbf{z}) = F_\eta(\Delta v(t,\mathbf{q},\mathbf{z})) - F_\eta(\Delta v(t^l,\mathbf{q},\mathbf{z})) \quad (11.)$$

This leads to the following extension of Hanemann *et al.*'s (1991) sample log-likelihood function:

$$\begin{gathered}\Sigma_j\Sigma_i I_{ji}^{1} I_{ji}^{2}\, ln[1\text{-}F_\eta(\Delta v(t^h,\mathbf{q}_{ji},\mathbf{z}_{ji}))] + \\ \Sigma_j\Sigma_i I_{ji}^{1}(1\text{-}\ I_{ji}^{2}) ln[F_\eta(\Delta v(t^h,\mathbf{q}_{ji},\mathbf{z}_{ji})) - F_\eta(\Delta v(t,\mathbf{q}_{ji},\mathbf{z}_{ji}))] + \\ \Sigma_j\Sigma_i(1\text{-}\ I_{ji}^{1})\ (1\text{-}\ I_{ji}^{2}) ln[F_\eta(\Delta v(t^l,\mathbf{q}_{ji},\mathbf{z}_{ji}))] + \\ \Sigma_j\Sigma_i(1\text{-}\ I_{ji}^{1}) I_{ji}^{2}\, ln[F_\eta(\Delta v(t,\mathbf{q}_{ji},\mathbf{z}_{ji})) - F_\eta(\Delta v(t,\mathbf{q}_{ji},\mathbf{z}_{ji}))],\ i = 1 \quad N,\ j = 1 \quad M. \quad (12.)\end{gathered}$$

where I_{ji}^{1} is the indicator function for a first positive response and I_{ji}^{2} is the indicator function for a second positive response.

To estimate these probabilities it is necessary to invoke distributional assumptions for η and a specification for Δv. For example, if both e_{ji} and e_{ji} are i.i.d. normal, their difference η is also distributed i.i.d. normal, leading to the probit model of Prob(Yes|t,**q**,**z**). If instead both e_{ji} and e_{ji} are i.i.d. extreme value type I, η is distributed i.i.d. logistically, giving rise to the logit model. Both the normal and the logistic distribution span all real numbers, while *WTP* for recreation for visitors interviewed at the forest site is unlikely to be negative. For this reason the distribution is often defined over a log transformation of the bid amount *t*. Furthermore, monotonicity of the log function preserves percentile estimates, such as the median, *M*(*WTP*). This property is useful to estimate the values of *WTP* for given percentiles of the population of visitors.

We propose that the increase in utility Δv derived from a visitor to a forest site depends on a vector of forest attributes **q** relevant to the outdoor experience as well as on one of individual characteristics **z,** and that this relation is linear with a given set of parameters {a, γ, b} to be estimated. Thus, Δv is specified as a linear index. So, assuming a linear index for $\Delta v = a + b\ln(t) + \gamma'\mathbf{x}$, where **x** collects the effects of **q** and **z**, and given that η is distributed logistically, one can re-write equation (5.) as:

$$\Pr(\text{Yes}|\mathbf{x},t;a,\gamma,b) = 1 - \Pr(\eta<\Delta v) = 1 - \Lambda(\Delta \text{v}) = \Lambda(-\Delta v) \qquad (13.)$$

and since $\Delta v = a + b\ln(t) + \gamma'\mathbf{x}$ we get:

$$\Pr(\text{Yes}|\mathbf{x},t;a,\gamma,b) = [1+\exp(-a - b\ln(t) - \gamma'\mathbf{x})]^{-1} \qquad (14.)$$

This is the log-logistic model in the presence of covariates **x**. In this specification the bid parameter b is the marginal utility of the natural log transform of money, γ is the vector of utility difference shifters associated with the covariates **x**, while a captures all the other effects in a constant. After estimating a*,* b and γ by maximum likelihood, the various features of the *WTP* distribution (expectation, median and other percentiles) can be computed from the parameter estimates. Under the correct specification, thanks to the invariance properties of maximum likelihood estimates (Goldberger, 1993), they will be unbiased and minimum variance estimates of the population parameters.

By definition, median *WTP* ($M(WTP)$) is the value of t at which Pr(Yes|Bid Amount) = 0.5. The logistic distribution is symmetric around zero, so $M(WTP)$ is the value of t such that $a+\gamma'\mathbf{x}+b\ln(t) = 0$, which leads to:

$$M(WTP|\mathbf{x},t;a,\gamma,b) = \exp[-(a+\gamma'\mathbf{x})/b] \qquad (15.)$$

It is also possible to obtain estimates of all the other percentiles as functions of the estimated parameters by using the equation:

$$WTP(p) = \exp[-(a+\gamma'\mathbf{x})/b + I_p] \qquad (16.)$$

Where I_p is the logistic variate corresponding to the chosen percentile p:

$$I_p = -\ln(1/p-1) \qquad (17.)$$

Equations (16.) and (17.) allow estimation of a probabilitistic demand function conditional on **x**. Computing p and $WTP(p)$ with and without a NR in a specific forest site, it is possibile to infer the changes in the *WTP* associated with NR creation.

However, the expectated value is sensitive to the log transform and by Jensen's inequality for a concave transformation $f(x)$, such as the log operation, $E[f(x)] > f[E(x)]$. In fact, in our particular specification it can be shown that (Duffield and Patterson, 1991):

$$E(WTP|\mathbf{x},t;a,\gamma,b) = \exp[-(a+\gamma\,\mathbf{x})/b]\,[\pi/b\,]/\sin[\pi/b] \qquad (18.)$$

This formula is employed to derive estimates of expected *WTP* from the ML estimates of the parameters a, γ, b. This calculation is not defined if the estimated b coefficient is higher than –1 (Duffield and Patterson, 1991; Ready and Hu, 1996). However, in this study we did not encounter this problem.

Survey administration and data

In 1992 the Queen's University of Belfast conducted a recreation benefit study by administering on-site face-to-face CV interviews in 13 sites in Northern Ireland and 13 in the Republic of Ireland. Over 9,400 visitors were interviewed by trained interviewers who completed the task in a period of a few weeks, short enough to ensure preference stability. All the CV surveys shared an identical design across forest sites he question asked to all respondents in all sites was:

> "If it were necessary to raise funds through an entry charge to ensure this forest or woodland remained open to the public and with no charge being made for parking, would you pay an entry charge of £ *t* for each person in your party (including young people under 18) rather than go without the experience?"

We are therefore comparing two states, the first in the *presence* of the outdoor visit to site j and the payment of the admission charge t which defines the state $u(m-WTP,f(\mathbf{q});\mathbf{z})$; the second, in the *absence* of the outdoor visit to site j and intact income level m, which defines the state $u(m;\mathbf{z})$. This money measure is an Hicksian compensating measure as it includes an income effect.

The inital (first bound) bid amounts t used were: {50, 100, 150, 250, 400} (in pence). They were uniformly distributed across visitors. Respondents who answered "yes" were presented with a follow-up question that reiterated the *WTP* question with a higher bid amount t^h respectively : {100, 150, 250, 400, 700}. Instead, respondents who answered "no" were asked the same question again, with a lower bid amount t^l respectively : {30, 60, 80, 150, 250}. Bid amounts were chosen on the basis of initial parameter estimates of the *WTP* distribution obtained from extensive pilot studies.

During the interview other information was also obtained concerning the socio-economic profile of visitors, such as age, sex, household income, personal income, dominant reason for the visit, means of transport to the forest and other information characterizing the profile of the visitor. All of these were included in the **z** vector. However, only household income had a statistically significant effect and was stable for different functional forms. This was hence combined with data on the site attributes deemed relevant for outdoor recreation, which made up the **q** vector. The forest attributes relevant for this paper are in Table 1.

The presence of a NR is a site-specific attribute and disentangling this effect from those of other attributes requires CV surveys be conducted across a number of sites, with and without a NR. There must be enough different sites to allow sufficient variation in site attributes to measure their effects on observed responses. Given the importance of bid design for welfare estimates the different sites should also share the same bid design. The Irish CV study has those desirable characteristics.

Estimating probability of response conditional on site attributes.

The Irish forest sites surveyed differed in many of the attributes that could affect a visitor's recreational experience. This study concentrates on a few that were measured, and that could be important for forest managers. The vector of site attributes **q** included total area (TOTAREA in 100 hectare units), under the hypothesis that the sheer extent of a forest could affect the experience of its visitors. It also included a dummy variable (NATRES= 1 or 0) to reflect the presence or absence of a NR in the forest, a major policy issue being the desirability of such reserves. To assess the impact of large old trees, which are such a salient feature of forest landscapes, the percent of total trees planted before the year 1940 was used (PRE1940). Another descriptor of the forest landscape included in **q**, was the percentage of land covered with conifers (CONIFS) broadleaves (BDLEAF) and larch (LARCH) (measured in ten percentage points to decrease numerical errors). A measure of site congestion (CONGEST expressed as 1,000 visits/car park space/year) was used to control for the negative effect of inadequate facilities and crowding on the utility of a visit.

The vector **z** consisted of only one variable: the visitor's household income bracket (HHINCOME) with expected positive sign, reflecting higher probability of a Yes response at a given bid *t* at higher income brackets. Other functional forms were investigated, but gave inferior log likelihood values.

Single bound (henceforth SB) and double bound (henceforth DB) parameter estimates were obtained by maximimizing the log-likelihood functions (7.) and (12.) respectively. Maximization was carried out with the Newton-Raphson algorithm and the standard convergence setting of the Gauss Maximum Likelihood package (Aptech Systems, 1997), using analytical gradient and Hessian. The results for the interviews conducted in the 26 forest sites are in Table 2. Despite missing data on household income there were still 8,371 usable observations. All the coefficient estimates have the expected sign in both SB and DB models. The values of the two likelihood functions at a maximum are not directly comparable as the DB includes a second set of responses, nor would be other conventional measures of fit based on likelihood values (such as the various pseudo R squares). The mean likelihood value ($\exp(\ln L/N)^{xk}$), where k=1 for SB and 2 for DB) is similar and quite high for both models, 0.62 for the SB model and 0.65 for the DB, indicating a good fit for both models, and a small improvement in fit by the latter.

In the SB model all but one parameter were significantly different from zero at the 1% or 5% level. The exception was the coefficient of the old trees (PRE1940), which was not significantly different from zero at conventional significance levels, but which was significant, with the expected positive sign, once the follow-up question was used in the DB estimation. Interestingly, given the purpose of this study, in both models the presence of a NR had a significant effect on the visitors' *WTP* for the recreational experience. The magnitudes of the coefficients of the tree coverage variables are in accord with expectations. Decreasing marginal utility implies that a marginal increase of the most common feature produces the lowest increase in utility. So conifers, which are widely represented across Irish forests have the lowest marginal effect, followed by broadleaves which are the second most common and then by larch - the rarest species in Irish forests - but quite colorful in autumn landscapes. As expected, income has a low but significantly positive effect in the probability of a positive response to any given bid amount. The total area of the forest site has a low positive effect.

Model applications

The models in Table 2 allow for a range of inferences useful for forest managers. Here, because of its higher precision, inference is carried out with the DB model, under the hypothesis that the assumptions invoked in model estimation are in fact true. Yet, any model is only a simplification of the economic reality and this needs to be born in mind in the interpretation of the inference results.

The model defines a mapping between per visit *WTP* and forest attributes. The percentiles of the *WTP* distribution in the population of visitors, rather than its expectation or median, may be more useful for forest managers. It can show the effect of varying the access charge on the number of visits, and thus on congestion. The same can be achieved by varying the level of an attribute, such as broadleaf coverage in favour of conifers for example.

Because of its relationship with the median voter attitude towards a public policy, several policy decisions may be judged by their estimated median *WTP* $M(WTP)$. The logit cdf is symmetric around zero, so the $M(WTP)$ is obtained when the value of the linear index is equal to zero (see equation (15.) above). For example, suppose there is an interest in establishing a new NR in one forest. Table 1a contains the attributes **q** for this site, from which one can build the index I^0 for the ex-ante condition with no NR:

$$I^0 = \alpha + \beta \ln(WTP) + \gamma' \mathbf{x}^0 \qquad (19.)$$

For the median $p = 0.5$ and by symmetry of the logit cdf $I^0 = 0$, so

$$I^0 = \alpha + \beta \ln(WTP_{p=0.5}) + \gamma' \mathbf{x}^0 = 0 \qquad (20.)$$

$$\rightarrow M(WTP^0) = \exp[-(\alpha + \gamma' \mathbf{x}^0)/\beta] \qquad (21.)$$

After the introduction of a NR, the vector of forest site attributes would change into the ex-post vector $\mathbf{x}^1$, hence

$$M(WTP^1) = \exp[-(\alpha + \gamma' \mathbf{x}^1)/\beta] \qquad (22.)$$

and the change in medianWTP for a single visit, due to the presence of a NR is:

$$\Delta M(WTP) = \exp[-(\alpha + \gamma' \mathbf{x}^0)/\beta] - \exp[-(\alpha + \gamma' \mathbf{x}^1)/\beta]. \qquad (23.)$$

If all the other forest attributes **q** and socio-economic variables **z** are unchanged, the ex-post attribute vector $\mathbf{x}^1$ will change by one element of **q** only: the dummy for the presence of the NR changing from 0 to 1.

Similarly for the expected value:

$$\Delta E(WTP) = \{\exp[-(\alpha + \gamma\, \mathbf{x}^0)/\beta] - \exp[-(\alpha + \gamma\, \mathbf{x}^1)/\beta]\}[\pi/\beta]/\sin[\pi/\beta] \qquad (24.)$$

Because all the *WTP* parameters of interest in this study are closed-form expressions of the ML parameter estimates, their confidence intervals were generated by sampling randomly 10,000 times from the multivariate normal distribution of the parameter estimates. This distribution is centered at the point estimates of $\{\alpha, \gamma, \beta\}$ and has a variance-covariance matrix $\mathbf{\Omega}$ approximated here by the inverse of the computed Hessian at the likelihood function maximum (Krinsky and Robb,1985).

Effects of creating NRs

The models for the *WTP* distribution for recreation in Irish forests developed above, are used to estimate the distribution of *WTP* for each forest *j*, conditional on the site characteristics of that forest, $\mathbf{q}_j$, and on the median income bracket of visitors at that site, $\mathbf{z}_j$. We compute these for all the forests currently without a NR. The estimates of expected and median *WTP* for a visit at each site, in the status quo conditions are in Table 3. These estimates are obtained exclusively on the basis of the forest's attributes and are obviously different from those based on models estimated only on the responses of visitors intercepted at each single forest site. On the other hand, the latter models would not be of use to conduct inferences based on forest site attributes, which is the main thrust of this study. Hillsborough and Douneraile show respectively the lowest and the largest *WTP* values. The median *WTP* ranges from 75 (±6) pence per visit at Hillsborough under the status quo, to 194 (±47) pence at Dourneraile, while mean *WTP* ranges from 102(±7) at the first site to a maximum of 262 (± 50) at the second. These values are similar to those reported in other British woodland studies (Willis, 1991; Willis and Benson, 1989).

The model is also employed to infer *changes* in *WTP* per visit associated with the creation of a NR at sites that did not have one. Figure 1 reports the full probability distribution of *WTP* before and after the introduction of a NR on Tollymore forest. Although, the accuracy of estimated percentiles decreases when moving away from the median, there is a clear separation of the distributions with and without a NR.

Table 4 shows the predicted change in *WTP* with creation of NRs in each of the forests currently without one. Again Hillsborough and Douneraile show respectively the lowest and the largest welfare changes. The median *WTP* would increase by as little as 16 (± 5) pence at Hillsborough, and as much as 41 (±14) pence at Douneraile, while mean

WTP would increase from a minimum of 22 (±6) pence at the first site to a maximum of 56 (±20) pence at the second.

If respondents had followed the rule in equation (3.) when answering the CV question, then changes in *WTP* measure welfare changes and, under the usual caveat (Broadway and Bruce, 1984), can be aggregated and used in benefit-cost analysis. We computed the yearly aggregate impact on visitors welfare from the introduction of NRs. This was done by multiplying the estimated per visit changes in *WTP* by the yearly number of visits to each forest. The results are in Table 5 and show that amongst Northern Irish forests, creating NRs at Tollymore and Hillsborough would increase welfare the most. NRs at Lough Key and Hazelwood would make the largest welfare contribution in the Republic of Ireland. The total yearly welfare increase due to creating NRs is estimated at £251,628 (£226,277-£278,718) in Northern Ireland and £318,042 (£265,103-£382,036) in the Republic of Ireland. However, these are probably lower bound estimates of the true changes in social welfare. In fact, respondents have revealed their *WTP* an access charge to visit the forest rather than going without the experience. So, other values associated with NR creation, such as increased property values in the forests' surroundings or existence value for habitat protection or creation are excluded from these estimates.

Conclusions

To estimate the effect of creating NRs in Irish forests we extend the classical random utility model intepretation of CV responses to account for forest attributes. We then estimate the parameters of this model to predict the probability of the *WTP* for a forest visit from a large scale CV survey across 26 forests. Both SB and DB estimates support the hypothesis that forest attributes are strong determinants of the utility of a visit.

In particular, the presence of a NR has a significant positive effect on the *WTP*. Other forest characteristics that influence *WTP* significantly are forest area, site congestion, number of old trees, and proportion of conifers, broadleaf species and larches (this least common species being most important). The models are applied to estimate the *WTP* distributions for each forest site as well as their mean and medians. We then predict the effects on the welfare of visitors from a policy that establishes new NRs in each forest currently without one. The total welfare change for the set of forests investigated here exceeds 570 thousand pounds per year. At the current frequency of forest visits this constitutes a considerable flow of benefits. A capitalization at a conservative discount rate of 3 percent gives a present value of approximately 19 million pounds. A more conservative figure would use the lower bound of the 90 percent confidence interval. This would still give a present value of welfare change from introducing NRs of approximately £7.5 million for Northern Ireland and 8.8 for the Republic of Ireland.

References

Alberini, A., 1995. Efficiency versus bias of WTP estimates: bivariate and interval-data models. Journal of Environmental Economics and Management 29:169-180.

Aptech Systems, 1997. Maximum Likelihood Reference Manual, Aptech Systems Inc. Maple Valley, Washington, USA.

Bishop, R. C. and T. A. Heberlein, 1979. Measuring Values of Extramarket Goods: Are Indirect Measures Biased?, American Journal of Agricultural Economics, 61(5):926-930.

Boxall P. C., W. L. Adamowicz, J. Swait, M. Williams, J. Louviere, 1996. A comparison of stated preferences methods for environmental valuation, Ecological Economics, 18:243-253.

Broadway R. W. and N. Bruce, 1984. Welfare Economics, Blackwell Publishers.

Cameron T. A., and J. Quiggin, 1994. Estimation using contingent valuation data from a dichotomous choice with follow-up questionnaire, Journal of Environmental Economics and Management, 27:218-234.

Cameron T. A., and M. D. James, 1987. Efficient Estimation Methods for “closed-ended” Contingent Valuation Surveys, Review of Economics and Statistics, May, LXIX(2):269-276.

Cameron T. A., 1988, A new paradigm for valuing non-market goods using referendum data, Journal of Environmental Economics and Management, 15:355-379.

Downing M. and T. J. Ozuna, 1996. Testing the reliability of the benefit function transfer approach, Journal of Environmental Economics and Management, 30:316-322.

Duffield J. W. and D. A. Patterson, 1991. Inference and optimal design for a welfare measure in dichotomous choice contingent valuation, Land Economics, 67(2)225-39.

Englin J. and R. Mendelsohn, 1991. A Hedonic Travel Cost Analysis for Valuation of Multiple Components of Site Quality: The Recreation Value of Forest Management, Journal of Environmental Economics and Management, November **21**(3):275-90.

Forest Service, Department of Agriculture for Northern Ireland, 1991. Conservation and the Forest Service HMSO, Belfast.

Goldberger A., 1993. A course in econometrics, Harvard University Press.

Hanemann M. W., and B. Kanninen, 1996. The Statistical Analysis of Discrete-Response CV data, Working Paper No. 798, Dpt. of Agr. & Res. Economics, Univ. of California at Berkeley.

Hanemann M. W., J. Loomis and B. Kanninen, 1991. Statistical Efficiency of Double Bounded Dichotomous Choice Contingent Valuation, American Journal of Agricultural Economics, November, 73(4):1255-1263.

Hanemann M. W., 1984. Welfare Evaluations in Contingent Valuations Experiments with Discrete Responses, American Journal of Agricultural Economics, 66 :332-341.

Hanemann M. W., 1989. Welfare Evaluations in Contingent Valuations Experiments with Discrete Response data: a reply, American Journal of Agricultural Economics, 71:1057-1061.

Hoehn J. P. and A. Randall, 1987. A Satisfactory Benefit-Cost Indicator from Contingent Valuation, Journal of Environmental Economics and Management, 14(3):226-47.

Krinsky I. e A. Robb, 1986. Approximating the Statistical Properties of Elasticities, Review of Economics and Statistics, 68: 715-719.

Langford I. H., I. J. Bateman, A. P. Jones, H. D. Langford, and G. Stavros, 1998. Improved estimation of willingness to pay in dichotomous choice contingent valuation studies, Land Economics, 74(1):65-75.

Lockwood M., J. Loomis and T. Delacy, 1992. A Contingent Valuation survey and benefit-cost analysis of forest preservation in East Gippsland, Journal of Environmental Management **38**:233-43;

Mattson L. and C-Z Li, 1995. How do different forest management practices affect the non-timber value of forests? An Economic analysis. Journal of Environmental Management **40**:79-88.

McConnell K., 1990. Models for referendum data: the structure of discrete choice models for contingent valuation, Journal of Environmental Economics and Management, January, 18:19-34.

McCurdy R. J., 1989. Nature Reserves in Northern Ireland Forests, HMSO, Belfast.

McFadden D., 1974. Conditional Logit Analysis of Qualitative Choice Behavior, in Zarembka, P. (editor), Frontiers in Econometrics IV, Academic Press, London/New York.

N.O.A.A., 1993. Natural resource damage assessments under the Oil Pollution Act of 1990, Federal Register, Vol. 58, pag.4601-4614.

Ready, R.C. and D. Hu, 1995. "Statistical approaches to the fat tail problem for dichotomous choice contingent valuation." Land Economics, November, 71(4):491-499.

Sellar C., J.-P. Chavas and R. J. Stoll, 1986. Specification in the logit model: the case of valuation of nonmarket goods, Journal of Environmental Economics and Management, 13:382-390.

Willis K., 1991. The Recreation Value of the Forest Commission Estate. Scottish Journal of Political Economy, **38**:58-75.

Willis K. and J. Benson, 1989. Recreational Values of Forests, Forestry, **62**:93-110.

Table 1a. Site attributes for Northern Ireland forests .

Forest site	Total area (100 of hectares)	Congestion (100 visits per car park space)	Natural Reserve	Trees before 1940 (% of total)	Tree coverage (% of total forest area)			Median Household income bracket*
					Conifers	Broadleaves	Larch	bracket
Tollymore	6.29	2.68	No	26	57	5	21	5
Castlewellan	6.41	1.38	No	12	44	7	17	5
Hillsborough	1.99	40.00	No	6	57	12	17	5
Belvoir	0.95	44.00	Yes	0	24	6	27	5
Gosford	2.51	1.39	No	2	40	21	0	4
Drum Manor	0.94	1.40	No	11	20	9	0	4
Gortin glen	14.60	1.17	No	3	70	2	3	4
Glenariff	11.82	1.75	Yes	2	67	1	7	5
Ballypatrick	14.61	0.85	No	0	81	0	3	4
Somerset	1.38	2.00	No	3	59	14	6	3
Florencecourt	13.93	0.50	Yes	1	32	5	0	5
Lough Navar	26.09	0.77	Yes	0	68	1	1	5
Castlearchdale	4.99	4.75	Yes	1	54	3	4	4

* Income bracket was 1 = under £3,999; 2 = £4000-£7,999; 3 = £8,000-£11,999; 4 = £12,000-15,999;   5 =16,000-19,999; 6 = 20,000-29,999;

Table 1b. Site attributes for Republic of Ireland forests.

Forest site	Total area (100 of hectares)	Congestion (100 visits per car park space)	Natural Reserve	Trees before 1940 (% of total)	Tree coverage (% of total forest area)			Median Household income bracket*
					Conifers	Broadleaves	Larch	
Lough Key	3.4	3.00	No	7.3	22	78	0	5
Hazelwood	0.7	20.00	No	0	7	93	0	6
Dun a Dee	2.4	5.00	No	2.6	51	48	1	6
John F Kennedy	2.52	1.70	No	0.4	35	60	5	5
Dun a Ree	2.29	3.00	No	2.2	64	36	0	6
Currachase	2	3.30	No	0.3	20	68	12	5
Cratloe	0.65	3.80	No	2.1	56	3	41	6
Douneraile	1.6	4.00	No	8.1	4	96	0	4
Farran	0.75	1.70	No	0.9	83	7	10	6
Guaghan Barra	1.4	5.00	No	4.2	46	12	42	6
Avondale	2.86	1.80	Yes	2.4	30	10	4	5
Killykeen	2.4	2.00	No	2.7	90	8	2	5
Glendalough	3.26	2.00	Yes	4.3	42	7	27	6

* Income bracket was 1 = under £3,999; 2 = £4000-£7,999; 3 = £8,000-£11,999; 4 = £12,000-15,999;   5 =16,000-19,999; 6 = 20,000-29,999;

Table 2. Parameters of probability of willingness to pay function.

Parameter	Single Bounded		Double Bounded	
	Ln=*L* – 0.4799		ln*L* = – 1.1267	
Constant	9.752	0.276***	10.633	0.221***
Forest area	0.018	0.007**	0.016	0.005***
Congestion♣	-0.358	0.029***	-0.358	0.023***
Natural reserve	0.581	0.067***	0.465	0.065***
Old trees	0.0007	0.0015	0.0025	0.001**
Coverage:				
Conifers	0.054	0.019***	0.057	0.017***
Broadleaves__	0.129	0.017***	0.113	0.014***
Larches	0.205	0.034***	0.130	0.028***
Income Bracket	0.082	0.016***	0.101	0.013***
Bid	-2.217	0.048***	-2.416	0.038***

***, **=significant at 1% and 5% level, respectively, for 8,371 observations.
LnL=log likelihood function.
♣ congestion values in Table 1 were scaled by 1/10.

Figure 1: Distribution of WTP for a visit at Tollymore forest, before and after the creation of a nature reserve.

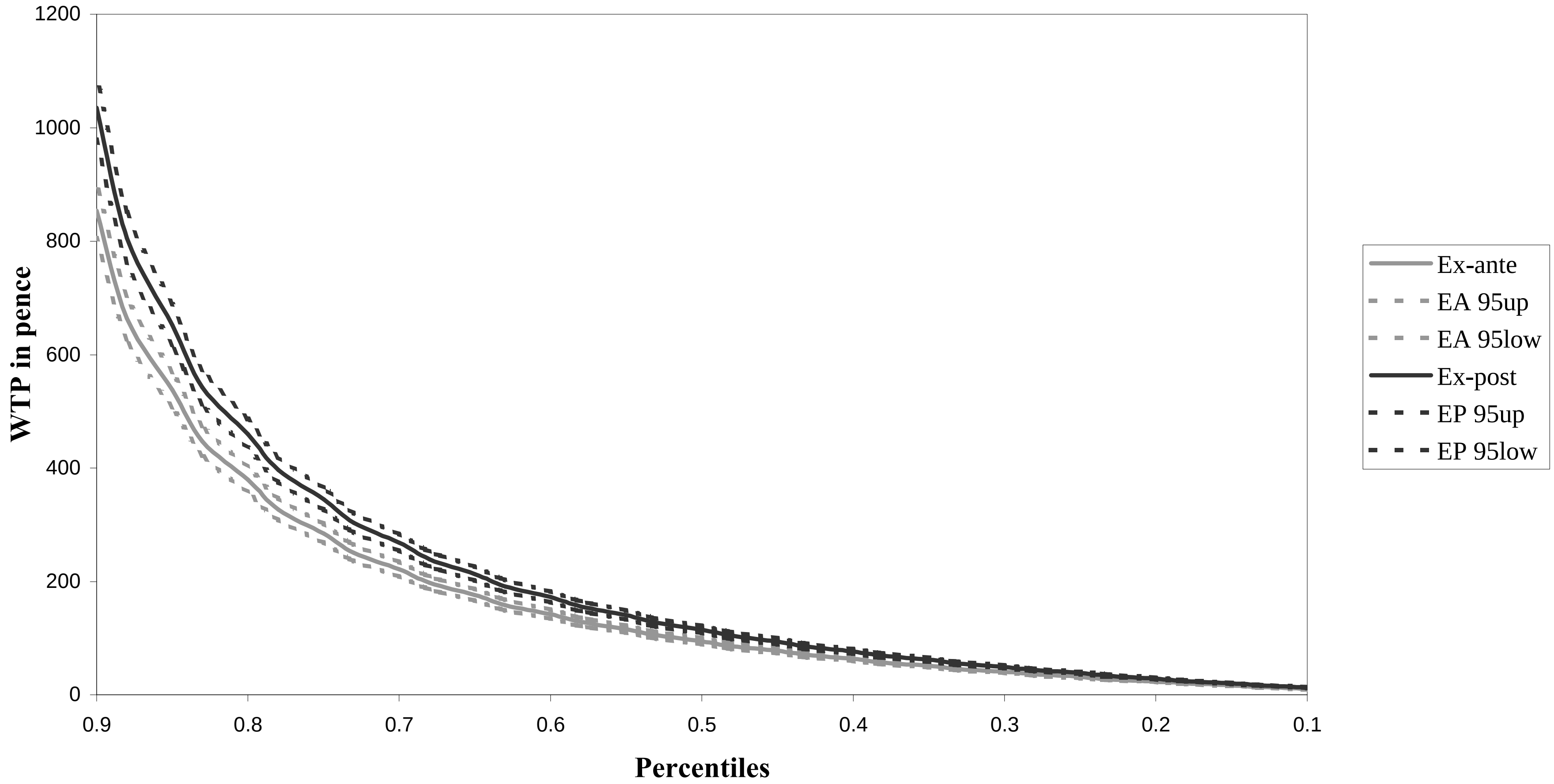

Table 3. Predicted WTP for a single visit in forests without a nature reserve.

Northern Ireland Forests							
Tollymore		Castlewellan		Hillsborough		Gosford	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
183 (175-192)*	136 (90-101)	175 (168-182)	129 (125-135)	102 (95-108)	75 (71-80)	160 (154-168)	119 (114-124)
Drum Manor		Gortin glen		Ballypatrick		Somerset	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
144 (135-154)	107 (100-114)	174 (166-184)	129 (123-136)	175 (166-186)	130 (123-137)	169 (162-176)	125 (121-130)
Republic of Ireland Forests							
Lough Key		Hazelwood		Dun a Dee		John F Kennedy	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
240 (204282)	178 (151-209)	214 (178-260)	159 (132-192)	191 (175-209)	142 (130-155)	221 (195-249)	163 (145-184)
Dun a Ree		Currachase		Cratloe		Douneraile	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
180 (168-194)	133 (124-143)	237 (205-274)	176 (152-203)	164 (154-174)	121 (114-129)	262 (212-326)	194 (157-241)
Farran		Guaghan Barra		Killykeen			
Mean	Median	Mean	Median	Mean	Median		
150 (140-162)	111 (103-120)	172 (164-181)	128 (122-134)	144 (133-157)	107 (99-116)		

***10% confidence interval.**

Table 4. Predicted changes in per visit WTP to forests without a nature reseserve, after creating one.

Northern Ireland Forests

Tollymore		Castlewellan		Hillsborough		Gosford	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
37 (27-48)	27 (20-35)	37 (27-48)	27 (20-35)	22 (15-28)	16 (11-21)	34 (24-44)	25 (18-33)

Drum Manor		Gortin Glen		Ballypatrick		Somerset	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
31 (22-39)	23 (17-29)	37 (27-47)	27 (20-35)	37 (27-48)	28 (20-35)	36 (26-47)	27 (19-34)

Republic of Ireland Forests

Lough Key		Hazelwood		Dun a Dee		John F Kennedy	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
51 (35-71)	38 (26-52)	46 (30-64)	34 (22-48)	41 (29-54)	30 (22-40)	47 (33-63)	35 (24-46)

Dun a Ree		Currachase		Cratloe		Douneraile	
Mean	Median	Mean	Median	Mean	Median	Mean	Median
38 (27-49)	28 (20-36)	50 (34-69)	37 (26-51)	35 (25-45)	26 (19-33)	56 (36-80)	41 (27-59)

Farran		Guaghan Barra		Killykeen	
Mean	Median	Mean	Median	Mean	Median
32 (23-41)	24 (17-30)	37 (26-47)	27 (20-35)	31 (22-39)	23 (17-29)

Table 5. Predicted welfare changes due to the introduction of a nature reserve, for the population of visitors at each site (Pound sterling per year).

Northern Ireland Forests			
Tollymore	Castlewellan	Hillsborough	Gosford
58,186	40,790	110,310	15,743
Drum Manor	Gortin glen	Ballypatrick	Somerset
7,109	11,081	5,656	2,743
Republic of Ireland Forests			
Lough Key	Hazelwood	Dun a Dee	John F Kennedy
76,515	45,510	40,610	40,291
Dun a Ree	Currachase	Cratloe	Douneraile
22,950	25,150	10,434	22,284
Farran	Guaghan Barra	Killykeen	
15,950	9,150	9,198	