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"The Value of Increasing the Length of Deer Season in Ohio"

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Abstract
Growing deer populations are controlled through changes in hunting regulations including changes in both hunter bag limits and season length. Such action results in direct benefits to hunters and indirect benefits to motorists and the agricultural sector as a lower deer population leads to fewer incidences of human-deer encounters. Traditional recreation demand models are often employed to examine the welfare implications of changes in daily hunting bag limits. Studies measuring the effects of changes in season length, however, are noticeably absent from the literature. This study uses a nested random utility model to examine hunter choice over site and season selection to derive the welfare implications of changes in season length.

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Introduction

In 1996 the State of Ohio generated nearly $13.7 million in revenue through the sale of licenses to deer hunters. In addition to generating revenue, selling licenses to hunt deer also creates a check on a thriving deer population estimated to be nearly 500,000. Indeed, such a healthy deer population imposes costs on society in terms of a high incidence of deer-vehicle collisions and large financial losses from crop damage. The Ohio Department of Public Safety has estimated that the approximately 26,000 deer vehicle collisions in 1996 alone resulted in losses of over $59 million dollars. Additionally, financial losses from crop damage due to deer nuisance can exceed $23 million annually (Forster and Hitzhusen, 1997). Alternatives confronting the Ohio Department of Natural Resources-Division of Wildlife for raising revenue while holding deer population in check include increasing the number of licenses or extending the deer-hunting season.

The purpose of this study is twofold. First, we estimate the per-trip willingness-to-pay of recreational hunters for an expansion of deer hunting season in twenty-three Ohio wildlife areas. Second, by modeling choice over both hunting site and hunting season, we illustrate a unique application of the traditional nested random utility model that permits an examination of both spatial and intertemporal choices.

Recreation demand models, which include travel cost models and discrete choice models, are commonly employed to value recreation related benefits (see, e.g., Bockstael, Hanneman, and Kling, 1980; Bockstael, McConnell, and Strand, 1989; Morey, Rowe, and Watson, 1993; Kaoru, Smith, and Liu, 1995). These models are based on the idea that the price of a recreational experience is represented at least in part by the costs incurred in accessing the recreation site. Random utility models of recreational hunting or fishing behavior model individual choices over sites and targets by specifying functions for the utility derived from the available alternatives. A trip utility function is specified to be a function of choice attributes, and it can be estimated via either a multinomial logit (MNL), nested multinomial logit (NMNL), or multinomial probit (MNP) model using data on individual trips and site characteristics.[1] For hunting, such attributes may include expected availability of a game species, on-site amenities, or limits on take. The estimated utility function can then be employed to measure the compensating variation from a policy change that is hypothesized to affect one or more of the site attributes.

Because of the inherent assumption that choices are made over a set of mutually exclusive alternatives, modelers are constrained to examine decisions made on a single trip occasion. The frequency of individual trips over a given season, though, can be estimated separately by specifying an aggregate trip demand function (see Parsons and Kealy, 1995, or Bockstael, Hanemann, and Kling, 1987). However, this type of model does not permit an estimation of the distribution and timing of trips within a given season, and thus it cannot be used to examine the welfare implications of changes in the length of the availability of a given target species over a broader or narrower season. By imposing the assumption that the trip choice under consideration is that of the single, non-repeated decision of when and where to take the first trip, the standard random utility framework can be extended to examine choices in an intertemporal context. Coupling this assumption with a non-overlapping set of season choices defined in terms of available game allows for an examination of the effects of changes in the season structure as well as for changes in the traditional site-specific factors.

The Nested Multinomial Logit Model of Hunter Season-Site Selection

A NMNL model is appropriate when modeling a number of discrete alternatives and similarities exist across the unobserved attributes of utility over particular choice decisions. In contrast to the multinomial logit specification (which imposes the independence of irrelevant alternative assumption (IIA) as illustrated in McFadden (1981)), the NMNL allows for the covariance across certain multidimensional choice

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[1] See McFadden (1989), Kling and Herriges (1995), Herriges and Kling (1996, 1997), and Morey (1999) for discussions on the MNL and NMNL models. McFadden (1989) and Herriges and Kling (1997) provide a good discussion of the MNP model.

decisions to be nonzero. In effect, the NMNL model acknowledges potential similarities across certain choice options confronting the agent. As such, the NMNL allows modelers the opportunity to group similar choice options thereby allowing for correlation patterns within groups to differ from correlation patterns across groups (Herriges and Kling 1996).[2] Another potential advantage of the NMNL model is that it can capture the sequential nature of a decision process if indeed one exists.[3]

In the present application modeling season-site choice, the hunter is assumed to make a sequential decision that includes first choosing the season and then conditional on choice of season, choosing a site. Hunter utility is assumed to be a function of characteristics surrounding the season-site choice with the added constraint that the choice among alternatives is mutually exclusive. It is assumed that the season and site combination that is chosen is the one that yields the maximum utility to the individual and is a function of each alternative's characteristics. These characteristics vary by season, by site, and jointly by site and season for any selected trip.

To formalize this model, assume that the k^{th} hunter's indirect utility from a visit during season *i* to site *s*, V_{sik}, is given by:

(1) $V_{sik} = V(Z_{sik}) + \varepsilon_{sik}$

Z_{sik} is a vector of observable (deterministic) variables that affect the utility derived from a visit to site *s* during season *i* by hunter *k*, and ε_{sik} is the random component of utility known only to the hunter.

Hunter *k* makes the decision to visit site *s* during season *i* only if:

(2) $V(Z_{sik}) + \varepsilon_{sik} > V(Z_{tjk}) + \varepsilon_{tjk}$ for all $j \neq i$, *and* $t \neq s$

Note that if the site and season decisions were considered independently, the vector Z_{sik} could be separated into attributes that influence each decision alone. However, for most hunters and trips, this is unlikely to be

[2] As Herriges and Kling (1996, 1997) suggest, while the NMNL is more flexible than the MNL (because the grouping allows for a relaxation of the IIA assumption), it still imposes more restrictions than the MNP model. That is, the NMNL still imposes certain restrictions on the variance-covariance matrix of disturbances terms while the MNP relaxes these restrictions (Ben-Akiva and Lerman, 1985). For a comparison of these three approaches, see Herriges and Kling (1997).
[3] Of course, if a sequential decision process does not seem justified, a simultaneous estimation process may be used (see Kling and Thomson, 1996).

the case. The set attributes that influence the first decision will likely include those that influence the latter choice. As we lack the information to discern which of the two decisions is made first, or if they are made jointly, we assume that hunters first choose a season for their trip, and then contingent upon that choice, select a site. Site characteristics, then, can be seen to influence both season and site choice.

With this structure, we can write a linear form for the indirect utility function in (1) as:

(3) $V_{sik} = \beta_1 x_{sk} + \beta_2 x_{ik} + \beta_3 x_{isk} + \varepsilon_{sik}$

Where x_{sk} is the set of attributes influencing utility that vary by site alone; x_{ik} is the set of attributes influencing utility that vary by season alone; and x_{sik} is the set of attributes influencing utility that vary by both site and season. The β's are vectors associated with these attributes.

Assuming that there are N_J potential sites and N_T seasons, the individual is seen as choosing between $(N_T \cdot N_J)$ mutually exclusive alternatives for their first trip.[4] Considering a structure that allows for correlation across sites that share the same season, but does not allow for correlation across seasons that share the same site, we can represent the joint probability of an individual choosing site i and season s as follows:

(4) $P(si) = P(s|i)\,P(i)$

where P(i) is the marginal probability of choosing season i and P(s|i) is the conditional probability of choosing site s given season i is chosen. If we assume that the error terms in (3) are independent and identically distributed as generalized extreme value random variables, the marginal and conditional probabilities that an individual will select a particular season *i* and site s conditional on season i, respectfully, are represented as:

(5) $$P(s|i) = \frac{\exp\left[\left(b_1 x_s + b_3 x_{is}\right)/(1-s_i)\right]}{\sum_{j\in N_J} \exp\left[\left(b_1 x_j + b_3 x_{ij}\right)/(1-s_i)\right]}$$

[4] Because some season-site combinations might not be selected, the actual number of feasible elements may be less than $(N_T \cdot N_J)$ when applied to a given set of data.

$$P(i) = \frac{\exp\left[\left(b_2 x_i + (1 - s_i) I_i\right)\right]}{\sum_{t \in N_T} \exp\left[\left(b_2 x_t + (1 - s_t) I_t\right)\right]}$$

where:

$$I_i = \ln \sum_{j \in N_J} \exp\left[\left(b_1 x_j + b_3 x_{ij}\right) / (1 - s_i)\right]$$

The term I_i is called the inclusive value for season i and contains information relevant to the season choice that is within the set of variables influencing the subsequent site decision. This term can be interpreted as the expected maximum utility from the set of site choices associated with a given season choice i. It captures information about the overall utility from the final season and site combination and hence influences the initial decision. The scale parameter, $1- \sigma_i$, measures the degree of substitutability across seasons with σ_i approximating the correlation coefficient across seasons.[5] Under the IIA assumptions from the multinomial logit, the value $(1- \sigma_i) = 1$ for all seasons. As noted in Kling and Thomson (1996), the recreational demand literature often assumes the scale parameter across all i's are equal (see, e.g., Morey, 1998).[6]

Under the sequential process assumed here, the conditional choice probabilities provide estimates of the $b_1/(1-\sigma_i)$ and $b_3/(1-\sigma_i)$ vectors to be used in estimating the inclusive values, which are treated as separate variables in the site choice stage. The scale parameter, $(1-\sigma_i)$, can then be estimated from the second stage by using the above parameter estimates. Once the scale parameter is estimated, the parameter estimates b_1 and b_3 can be recovered. Substituting the equations in (5) into equation (4), the joint probability that an individual will select site *s* and season i on a given trip occasion is estimated as:

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[5] McFadden has suggested that while σ is not equal to the correlation coefficient it is a close approximation such that: $\sigma \leq \rho \leq \sigma + 0.045$. (Maddala, 1983: page 71)

[6] Furthermore, for NMNL models to be globally consistent with utility maximization, $1-\sigma_i$ must lie between zero and one for all i (Daly and Zachary, 1979; McFadden, 1981). Borsch-Supran (1990) and Herriges and Kling (1996) have provided conditions for which the NMNL models may be locally consistent with utility maximization when the scale parameter is greater than one.

$$(6) \qquad P(si) = \frac{\exp\left[\left(b_1 x_s + b_3 x_{is}\right)/(1-s_i)\right]}{\sum_{j\in N_J} \exp\left[\left(b_1 x_j + b_3 x_{ij}\right)/(1-s_i)\right]} \; \frac{\exp\left[\left(b_2 x_i + (1-s_i) I_i\right)\right]}{\sum_{t\in N_T} \exp\left[\left(b_2 x_t + (1-s_t) I_t\right)\right]}$$

where I is defined above. Graphically, such a structure could be represented as follows:

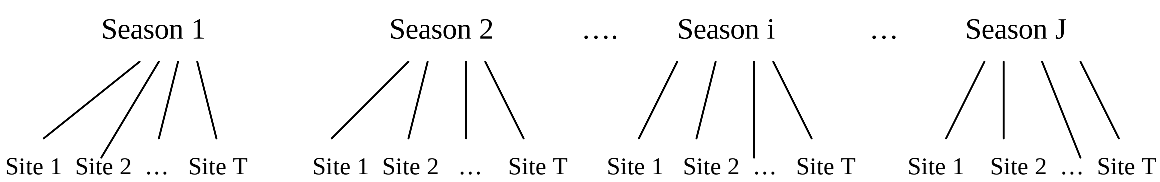

Figure 1. Nesting Structure under Sequential Decision Process for Season-site Choice

Given that we intend to look at the effects of increasing season length on hunter utility, an issue arises with how the correlation among disturbances changes both within and across groups (seasons) with a marginal change in one of the seasons. As mentioned above, it is not uncommon to assume $\sigma_i = \sigma \ \forall$ i. In our application, a marginal increase in season length has a modest impact, if any, on the characteristics that vary only across seasons. Thus, we assume that the cov (U_{is}, $U_{is'}$) = var (ε_i) for season i and sites s and s' is approximately equal to the cov ($U_{i's}$, $U_{i's'}$) = var($\varepsilon_{i'}$) for the new season i' and sites s and s'. Furthermore, since where assume the scale effect is constant across seasons, var (ε_i) ≈ var ($\varepsilon_{i'}$) ≈ var (ε_j), where j ≠ i ≠ i', and $\varepsilon_{(.)}$ is generated by an extreme value distribution with scale parameter (1- σ) and a location parameter η_i that we assume to be zero for all i.

Data

This study relies on data from a survey of recreational activities in Ohio conducted by the Institute for Local Government and Rural Development (ILGARD) at Ohio University. The intercept survey was distributed over thirteen months beginning in April 1996 and targeted twenty-three of Ohio's seventy-nine wildlife areas (sites) using stratified random sampling according to the regional districts in which they were located. Ohio is comprised of five regional districts (see Map 1). For the survey, four sites were selected from Districts One and Five, and five sites from the remaining three districts. In selecting these twenty-three sites, the criteria included allowing for sites that varied in size, use, and habitat. All sites selected were owned by the state for at least two years and were not designated as wildlife protection areas.

A one-page survey was placed on the windshield of the cars parked within the wildlife areas and included questions regarding types of activities conducted, frequency of visits, trip duration, number of people in the party, where the party was staying, trip expenditures, why this area was chosen, distance traveled, and demographics. Overall, 14,627 surveys were distributed from which 5,015 were returned, resulting in a 34.3% return rate. Of the 5,015 respondents, 4,870 (or 97.9%) indicated they resided in Ohio. The distribution of visitor's residences was quite large, with all but a single county in the state represented.

Of those surveyed, 47.3% (3,157) specified hunting as their purpose for visiting the area, 25.9% (1,727) specified fishing, 6.9% (462) suggested bird watching, while 19.9% (1,362) came for other activities (e.g., target shooting, hiking, dog training, boating, and camping). The top five hunting activities included: deer (1513 visitors), pheasant (1,042 visitors), rabbit (1,004 visitors), squirrel (997 visitors), and turkey (697 visitors). Other hunting activities included targeting woodcock, quail, duck, grouse, geese, dove, fur-bearers, and other small game.

A majority of respondents indicated they had visited the site in the past twelve months, however, no information was collected regarding whether those trips were taken during the fall or spring seasons. Seventy-five percent of those interviewed suggested staying at the site between one and two days. Ninetysix percent indicated they were planning on returning to the site within the next twelve months. Approximately eighty percent indicated they visited other wildlife areas over the last twelve months and sixty-eight percent had visited a wildlife area within fifty miles of the one they were visiting. In traveling to the wildlife areas, around one-third of all visitors traveled between twenty-one and fifty miles one way, while forty percent traveled less than 20 miles one way. When asked if they had completed this survey before, 87.7% responded no. Finally, 33.6% of the respondents indicated a salary of $50,000 or more, and 25.9% listed an income between $35,000 and $49,999.

The large number of respondents and the detail associated with the questionnaire allows for the derivation of compensating variation values for changes in species-specific season length using a nested random utility model of site and season choice. We focus our attention on those hunters engaged in a single-day hunting experience, of which there were 983 respondents. We also target the Fall 1996 hunting season, from September 5 through December 31. This period of time is divided into four non-overlapping "seasons", which are defined according to legally available game and allowable weapons, as table 1 illustrates.[7]

Table 1. Ohio Hunting "Seasons" (1996)

Season	Dates	Game Available	Length in days
1	9/5-10/18	Squirrel, Dove, Grouse	44
2	10/19-12/1	Squirrel, Dove, Grouse, Waterfowl, Turkey, Pheasant, Quail, Fox, Raccoon, Opossum, Deer (archery)	44
3	12/2-12/14	Deer (gun and archery), Waterfowl	13
4	12/15-12/31	Squirrel, Dove, Grouse, Waterfowl, Pheasant, Quail, Fox, Raccoon, Opossum, Deer (archery), Deer (primitive weapon)	17

[7] Strictly speaking, each day is a viable choice alternative. When the time period of concern is divided into game-specific sections, there are 14 "seasons". As our focus is the deer (gun) season, we aggregate both before and after this season in order to simplify the modeling task.

Methodology

To estimate the model outlined above, we account for variables that vary over season alone, site alone, and jointly over season/site. We specify that season choice is a function of available game, weather, and season length. We construct a dummy variable for season 3, to capture the effect of removing the small game species and introducing the deer (gun) as a viable target. As colder weather may be a deterrent to any type of recreation, we use the average daily low temperature during the season as a quality variable in the season choice stage. The length of a season may also influence season choice. That is, a longer season increases the opportunities to select that season for a trip. To account for this, and to compensate for potential bias due to aggregation of the smaller choice alternatives (seasons or days), we follow Ben-Akiva and Lerman (1985) and use the log of the number of days as a quality variable in the season choice stage of the indirect utility function.

Site choice may be influenced by many factors. The 23 sites in our survey differ with regard to many attributes such as size, presence of water bodies, shooting/target ranges, and areas for dog training. We use the acreage of each site to capture the effect of size on site choice, and construct dummy variables for sites with large water bodies, shooting ranges and dog trial areas. Travel costs, which vary across sites but not season, are estimated as the sum of explicit travel costs (estimated at $0.30 per mile) and the opportunity cost of travel time.[8] We include two variables that vary across both site and season -- bag limit and a quality variable indicating availability of deer. Five sites in our study have a one-deer limit during the deer season. We attempt to capture this effect by using a dummy variable for those sites. Regardless of season, some sites may be more favorable for hunting deer. To capture the effect of the historical availability of deer during seasons when deer are a viable target, we use the number of deer harvested in the

[8] Driving time was calculated assuming 45 miles per hour average speed. The software package HYWAYS/BYWAYS was used to generate travel distances. Two-thirds of the wage was used as an approximation of the opportunity cost of time. Hourly wage was estimated by dividing reported income by 2080. A simple OLS regression with estimated miles (using the program HYWAYS/BYWAYS) as the dependent variable and miles reported by the respondents as the independent variable revealed an intercept of 31 and a parameter estimate of 0.79, with a p-value of .0001.

previous year (1995) in the county of each site as a variable influencing site choice.[9] Table 2 provides a description of variables used in the model.

We can now write an explicit form for equation (3), the indirect utility from a hunting trip during season *i* by a hunter to site *s*, as:

(7) $V_{si} = \beta_1(\text{travel cost}_s) + \beta_2(\text{area}_s) + \beta_3(\text{past deer harvest}_{si}) + \beta_4(\text{deer bag limit}_{si}) + \beta_5(\text{dog area}_s) + \beta_6(\text{water}_s) + \beta_7(\text{shooting range}_s) + \beta_8(\text{low temperature}_i) + \beta_9(\text{log of number of days}_i) + \beta_{10}(\text{deer dummy}_i) + \beta_{11}(\text{inclusive value}_s)$

Table 2. Variable descriptions

VARIABLE NAME	DESCRIPTION	VARIATION
Cost	Travel cost to site	Varies over sites
Acres	Total acres of site	Varies over sites
PDH	1995 Deer harvest at site	Varies over sites for seasons 2,3,4 and = 0 for season 1
Limit	Dummy = 1 if one-deer limit at site, = 0 otherwise	Varies over sites for seasons 2,3,4 and = 0 for season 1
Ran	Dummy = 1 if shooting range at site, = 0 otherwise	Varies over sites
Dog	Dummy = 1 if dog trial area at site, = 0 otherwise	Varies over sites
Lake	Dummy = 1 if lake at site, = 0 otherwise	Varies over sites
Lot	Season's average low temperature	Varies over seasons
Lnd	Log of number of days in the season	Varies over seasons
Deer	Dummy = 1 if deer gun season, = 0 otherwise	Varies over seasons
I	Inclusive value from site choice stage	

[9] Past year's deer harvest were not available by wildlife area. When a wildlife area bordered two or more counties, the average harvest for the counties were used.

Results

We estimate (7) using the sample of 983 single-day hunters for the fall season of 1996. The estimated coefficients and standard errors are reported in Table 3. These coefficients can be interpreted as the marginal trip utilities for the corresponding variables. A positive coefficient is therefore an indication that the variable influences the site or season choice in a positive fashion. We see that sites with higher travel costs contribute less to utility and are therefore less likely to be selected. Larger sites may offer more or better hunting opportunities than smaller sites, as sites with a greater number of acres are more likely to be chosen. The availability of deer at a site as indicated by the previous year's harvest appears to strongly influence site choice. The positive sign on the one-deer limit coefficient is somewhat puzzling, as we would expect that hunters would be averse to sites where there is a bag limit. Perhaps individual expectations do not exceed a single deer per season; hence the one-deer limit does not affect site choice in a negative fashion. There may also be some quality difference at these sites, perhaps the size of deer or availability of other game, which we have not captured with our variables. Sites with target/shooting ranges are more likely to be selected, while site with dog training areas or water bodies are less likely to be selected. These results may follow from the fact that our sample is limited to hunting trips. As such, target practice may be viewed as a complement to a hunting experience. Dog training, meanwhile, may be thought of as leading to a decrease in the presence of wildlife in the surrounding vicinity. Sites with significant water bodies may be primarily used for other types of recreation, such as boating, which again may decrease the presence of wildlife. Another potential explanation for the negative sign on water bodies is that since deer hunters make up a significant portion of our sample, and optimal deer habitat is characterized by either the presence of brushland or woodland.

The season choice variables are all of the expected sign. Lower temperature does appear to discourage hunting activities, and longer seasons makes it more likely that a trip will be taken. Clearly, hunting deer with guns is a preferred activity, given the sign and magnitude of the deer season coefficient. Notice also that the coefficient on the inclusive value in the second choice stage is 0.365, yielding an estimate of 0.635

Table 3. Indirect Utility Function Coefficients

VARIABLE NAME	COEFFICIENT	STANDARD ERROR
Cost	-.044085 ***	.001401
Acres	.000056 ***	.000009
PDH	.000638 ***	.000057
Limit	.560161 ***	.137746
Ran	.491648 ***	.127244
Dog	-.166628 **	.076306
Lake	-.250960 ***	.087897
Lot	-.046037 ***	.005403
Lnd	3.487066 ***	.245186
Deer	2.112459 ***	.355213
1 - σ	.365235 ***	.109490

*** significant at 1 percent level, ** significant at 5 percent level, * significant at 10 percent level.

for σ, which is significantly different than both zero and one. This suggests that a non-nested model would likely result in a violation of the assumption of independence of irrelevant alternatives.

As travel costs enters the indirect utility function in a linear fashion, the coefficient on travel costs in equation (7) can be considered to represent the marginal utility of income. Given our estimate of (7), the benefits of an improvement in the quality of one of the site or target characteristics can be estimated as the per trip compensating variation for the nested logit model:

$$\text{(8)}\quad CV_k = \frac{1}{\beta_1}\left\{ \ln \sum_{i=1}^{N_I}\left[\left(\sum_{s=1}^{N_T} \exp\left[\frac{V_{is}^{0}}{(1-\sigma)}\right]\right)^{(1-\sigma)}\right] - \ln \sum_{i=1}^{N_I}\left[\left(\sum_{s=1}^{N_T} \exp\left[\frac{V_{is}^{1}}{(1-\sigma)}\right]\right)^{(1-\sigma)}\right]\right\}$$

where V^0 and V^1 represent utility before and after the change in season.

In order to derive the willingness to pay for an extension of the deer (gun) season, we estimate the per-hunter compensating variation for a one-day increase in the number of days in season 3 by increasing the log of the number of days choice variable for that season. We calculate this measure first without simultaneously altering the length of the other seasons. However, as deer season is clearly a substitute for the alternative seasons, it makes more sense to shorten one of the adjacent seasons while lengthening season 3.[10] The extra day for deer hunting can therefore be taken from season 2 or season 4. The net benefits to existing hunters then would be the combination of losses from shortening season 2 or season 4 and benefits from lengthening season 3.[11]

The mean per-hunter compensating variation for a one-day increase in deer season are reported in Table 4 below. We can interpret these values as the mean willingness to pay per hunter for the option of having a 14-day deer-hunting season instead of a 13-day season. Notice that the mean willingness-to-pay measures are higher when the extra day is taken from season 4. This is due likely to the fact that season 4 is a less popular alternative. Thus, shortening this season is less costly in terms of utility lost than shortening season 2.[12]

Assuming that the hunters in our sample are representative of the larger Ohio deer hunting population, we can further expand the values in table 4 into state-wide welfare impacts by multiplying the number of deer hunters in Ohio by our mean willingness-to-pay estimates. In 1996 approximately 404,141 deer permits were sold. Multiplying the mean willingness-to-pay per hunter by 404,141 yields a total willingness

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[10] Notice in Table 1 that the Ohio Department of Natural Resources limits the availability of other species during deer (gun) season.

[11] We do not account for changes in participation rates for existing hunters, nor do we allow for new entrants following the expansion of season length. Given that one season is expanded while another is shortened, we cannot state a priori if there would be a net increase or decrease in participation. These changes could be estimated and incorporated into the analysis using a linked model (see, e.g., Bockstael, Hanneman, and Strand, 1986; Bockstael, Hanneman, and Kling, 1987; Hausman, Leonard, and McFadden, 1995; or Herriges, Kling, and Phaneuf, 1999), or the repeated nested logit (see, e.g., Morey, Rowe, and Watson, 1993). This type of application is beyond the scope of this paper.

[12] This hypothesis is supported by the fact that the effects on the probabilities of taking the trip during seasons where length does not change are greater when the extra day is taken from season 2. That is, when season 2 is shortened, the decrease in the probabilities of taking the trip during seasons 1 and 4 is larger than the decrease in the probabilities of taking the trip during seasons 1 and 2 when season 4 is shortened. This indicates greater substitution of season 4 trips for season 3 trips than the substitution between seasons 2 trips for season 3 trips.

Table 4. Per Hunter Compensating Variation for One-Day Increase in Season 3 Length

Season Length change	Mean per-hunter compensating variation (standard deviation)
One-day increase in season 3 with no decreases in other season lengths.	\$3.41 (1.67)
One-day increase in season 3 with one-day decrease in length of season 2.	\$2.92 (2.02)
One-day increase in season 3 with one-day decrease in length of season 4.	\$3.36 (1.69)

to pay for a one-day extension of the deer (gun) season of approximately \$1.18 million if the day is subtracted from season 2 and \$1.36 million if the day is subtracted from season 4.

Discussion

A practical application of these results is to aid the Ohio Department of Natural Resources in meeting its stated objective of maximizing recreational benefits while minimizing the costs associated with harmful deer-human interaction. There are a number of deer management options available to the Ohio Department of Natural Resources in managing its deer population, including increasing the number of permits to hunt deer, changing the doe-to-buck permit ratio, or extending the deer season. While comparisons of these alternative policy options are sparse, we provide a modest attempt to analyze the latter of these options -- the welfare implications of extending season length. Our results suggest that the seasonal willingness to pay for a one-day extension of the deer season is between \$1.18 and \$1.36 million. Considering the indirect effects of decreasing deer population from an extension of the season that include both a potential reduction in deer-vehicle collisions and deer-related crop damage, the expected welfare implications from extending deer season would likely be even larger than the estimates provided above.

From a theoretical perspective, this application has given us a unique opportunity to focus on the welfare implications of an intertemporal choice. Traditional random utility applications have tended to highlight the welfare implications of changes in bag (or catch) limits or quality characteristics using site choice models. This analysis increases the choice set available to policy makers thereby providing a more realistic and accurate approach to modeling the implications of policy choice associated with wildlife management decisions.

This modest extension to the nested logit model literature is intended to give modelers more flexibility in analyzing potential policy options. Clearly, the results are a function of the restrictions imposed such as assuming a sequential decision process, the equality of the scale parameter (dissimilarity coefficient) across seasons, and a one-time trip scenario. While our results suggest globally consistent estimates, the restrictions we impose can influence the magnitude of our welfare measure. Incorporating sensitivity analyses that focuses on these restrictions and linking the season/site model to a participation model will help to broaden this usefulness of these results in the future.

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