Appendix B: Willingness-to-Pay for Bobcat Harvest Permits in Indiana

**B.1 The Turnbull Estimator**

Following Haab and McConnell (2002), we estimate WTP using the Turnbull estimator in two steps. In the first step, we estimate the distribution of “no” responses—i.e., the distribution of respondents who are offered a license at a particular price but indicate they would not buy it. Denote the bobcat license prices that the respondents were shown as *tj*, *j* = 1, …, *M*, with *tj* > *tj*–1 ∀*j*. Let the number of respondents who were shown *tj* and were willing to pay that amount be *Yj*. Likewise, let the number of respondents who were shown *tj* and were not willing to pay that amount be *Nj* such that the total number of respondents shown price *j* is *Tj* = *Yj* + *Nj*. Finally, define the share of respondents who said they would not purchase the license at price *j* as *Fj*. We can estimate the distribution of responses *Fj* ∀*j* by maximum likelihood. Given that individual responses are independent, and respondents are randomly assigned to prices, we can model the probability of observing *Yj* from a sample of *Tj* respondents using a binomial density with success probability 1 – *Fj*. The likelihood contribution for the subsample of respondents shown price *j* is then

(B1)

Define *fj* = *Fj* – *Fj* – 1 as the weight of the distribution that falls between prices *j* and *j* – 1. Assume also that the response data is appropriately pooled such that *Fj* + 1 > *Fj* ∀*j*—i.e., the

probability of a “no” response increases monotonically with price (c.f. Haab and McConnell 2002).[[1]](#footnote-1) We can then write the log-likelihood maximization problem as

(B2)

note that the term from (B1) does not depend on the distribution parameters *Fj* and hence can be ignored in (B2). The maximum likelihood estimators are

(B3)

The are the density functions for our response data.

 Once the density is estimated, we can calculate a lower-bound estimate of mean WTP for a bobcat license as

(B4)

 where the bid rejection proportion at bid *M* is 1. Additionally, we can calculate 95% confidence intervals for our WTP estimation. The variance of WTP is

(B5)

where *F̂j* = *Nj*/*Tj*. The 95% confidence interval is then calculated as

(B6) ,

where 1.96 is the critical value for a two-tailed normal distribution with infinite degrees of freedom.[[2]](#footnote-2)

*B.1.1 Willingness-to-Pay Estimates*

We use the Turnbull estimator to estimate WTP using different subsamples of our data. The first bar in Figure B.1 shows WTP for the baseline bobcat license (i.e., a bag limit of one animal and a state-wide quota of 300) among all 101 respondents who were shown the baseline combination of license attributes, regardless of their answers to the consequentiality questions in Figure 3 in the main text. We estimate a mean WTP of $21.73, with a 95% confidence interval between $15.85 and $27.60.

When restricting our sample to only those respondents who agreed with the consequentiality questions in Figure 3, WTP increased to $27.80. This estimate is statistically significantly different from both 0 and mean WTP among those who do not agree with any of the consequentiality statements at the 95 percent level ($3.99; see Figure B.1).

The reduction in mean WTP from our consequentiality test in the Turnbull estimates could also be due to pooling. For instance, the subsample of individuals receiving the bag limit = 1, quota = 300 combination that do not agree with any of the consequentiality statements are pooled at higher license prices, which could produce a lower WTP estimate. However, when we employ common pooling for those that agree and disagree, the result still holds. Further investigation reveals that individuals who do not agree with any consequentiality statement have a greater proportion of individuals responding “no” to our WTP question compared to individuals that agree. Therefore, the proportion of respondents indicating they would not purchase the license, *Fj*, is greater for those that do not agree, reducing WTP.

As in the main text, we also test for scope effects by comparing estimates of WTP across the subsamples of respondents shown different combinations of the bobcat bag limit and state-wide quotas. Testing for scope effects using the Turnbull estimator is complicated by two features of our data. First, each subsample required different pooling to ensure that the probability of “no” responses increases monotonically with price (i.e., that ; see footnote 4). Second, the prices we showed to respondents were not evenly spaced; the difference increased for higher prices. As a result, pooling on different prices can have a disproportionate impact on mean WTP estimated across different subsamples (see equation (B4)), masking potential scope effects.

We overcome this by conducting pairwise comparisons of different bag limit and quota combinations in which we pool our data on common prices for each comparison. We conduct four comparisons:

1. bag limit = 1, quota = 300 versus bag limit = 1, quota = 600;
2. bag limit = 1, quota = 300 versus bag limit = 2, quota = 300;
3. bag limit = 1, quota = 600 versus bag limit = 2, quota = 600; and
4. bag limit = 2, quota = 300 versus bag limit = 2, quota = 600.

The bars labeled “scope effect comparisons” in Figure B1 present the results of our scope effects analysis for the Turnbull estimator. When holding the bag limit constant at one animal, mean WTP decreases from $11.86 to $9.29 as the quota increases. However, the difference is small and not statistically significant. Similarly, when holding the quota constant at 300, mean WTP decreases from $20.39 to $14.58 as the bag limit increases. The decrease in mean WTP is relatively large in magnitude, though it is again not statistically different from zero. In contrast, when holding the bag limit constant at twoanimals, increasing the quota increases mean WTP by nearly 60%, although the difference is again statistically insignificant. Similarly, increasing the bag limit holding the quota fixed at 600 animals again increases mean WTP from $8.63 to $13.12. This difference is statistically significant at the 10% level.

We can motivate these findings by appealing to the intuition from our theoretical model derived in the main text. If the quota binds, then increasing the quota will increase the marginal individual’s expected harvest and, hence, economic welfare—and likewise for the bag limit. This is consistent with prior studies of Norwegian grouse hunters (Wam et al., 2012) and Louisiana waterfowl hunters (Gan and Luzar 1993) which find that hunters’ WTP for a permit increases with the daily bag limit. Nilsen et al. (2012) find that quotas for Norwegian lynx harvests are often nearly or completely filled and that harvest numbers increase with the quota, suggesting that WTP should rise as well. However, if harvesters perceive the quota to be binding, we might expect that increasing the bag limit would decrease a harvester’s WTP due to perceived congestion or competition for harvests. If the bag limit is binding, then increasing the quota may have no effect on WTP since mean harvest per individual will not increase.

These explanations are consistent with the scope effects we estimate in Figure B1, and our survey data provides some weak supporting evidence of this. In our survey, we asked respondents who were unwilling to purchase the bobcat license to indicate why they would not do so (see question 1.1.1., Figure 2). Depending on the combination of license attributes, 34–48% of respondents indicated that the price they were shown was too high. But among those that did not object to the price, Table B1 shows that none of the respondents shown licenses with a bag limit of one considered the quota to be binding, regardless of the quota. However, ~18% of those shown the lower quota and bag limit reported the bag limit to be too low. This suggests the bag limit is binding to these respondents and, hence WTP should be nonresponsive to changes in the quota. This is consistent with Figure B1, which shows that the estimated mean WTP is very similar under both quotas when the bag limit is one animal. Our Turnbull estimates also show that mean WTP decreases with the bag limit when holding the quota fixed at 300 (albeit not significantly). Table B1 shows that the percentage of non-buyers reporting the quota to be too low increases from 0 when the bag limit is one animal to 5.3% when the bag limit increases to two animals. This provides some evidence that these respondents consider the quota to be binding, which makes the decrease in mean WTP reported in Figure B1 consistent with our intuition: WTP goes down because harvesters’ expectations of harvesting a bobcat decrease, as the bag limit makes it easier (for others) to fill the quota. Finally, relaxing the binding quota (i.e., increasing the quota from 300 to 600 with the bag limit of two animals) or relaxing the binding bag limit (increasing the bag limit from one to two animals at a quota of 600) increases the estimated mean WTP in Figure B1. Only the latter effect holds in our parametric model. Table B1 suggests that respondents perceive neither the bag limit nor quota to be binding in these cases.

**References**

Haab, T.C. and K.E. McConnell. 2002. *Valuing Environmental and Natural Resources*. Edward Elgar: Northampton, MA, p. 70.

Whitehead, J.C. 2017. “Who Knows what Willingness to Pay Lurks in the Hearts of Men? A Rejoinder to Egan, Corrigan, and Dwyer.” *Econ Journal Watch* 14(3):346–361.

**Table B1** Shares of respondents unwilling to purchase license due to restricted harvest limits

|  |  |
| --- | --- |
|  | Share of non-buyers unwilling to purchase the bobcat license because: |
| License attributes | The bag limit is too low. | The quota is too low. |
| Quota = 300, bag limit = 1 | 17.65 | 0 |
| Quota = 300, bag limit = 2 | 10.53 | 5.26 |
| Quota = 600, bag limit = 1 | 0 | 0 |
| Quota = 600, bag limit = 2 | 0 | 0 |

**Figure B1** Willingness-to-pay estimates for the Turnbull model

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1. If *Fj* + 1 < *Fj*, then responses are pooled such that $T\_{j}^{\*}$ *= Tj + Tj+1*, with $Y\_{j}^{\*}$ and $N\_{j}^{\*}$ pooled similarly. Then the updated *Fj*, now denoted as $F\_{j}^{\*}$, is calculated as above. To demonstrate, in the following table we see that *F*9.25 *< F*6.75, indicating pooling is necessary. We then combine *Tj, Nj,* and *Yj* for *j =* 6.75, and *j =* 9.25, to create $T\_{j}^{\*}$, $Y\_{j}^{\*}$, and $N\_{j}^{\*}$. Note that $F\_{j}^{\*}$ is monotonically increasing, as expected.

|  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- |
| Price | Tj | Nj | Yj | Fj | Tj\* | Nj\* | Yj\* | Fj\* |
| 6.75 | 11 | 5 | 6 | 0.4545 | 21 | 8 | 13 | 0.3810 |
| 9.25 | 10 | 3 | 7 | 0.3000 | Pooled back |
| 11.25 | 38 | 17 | 21 | 0.4474 | 38 | 17 | 21 | 0.4474 |
| 19.25 | 17 | 9 | 8 | 0.5294 | 17 | 9 | 8 | 0.5294 |
| 24 | 9 | 5 | 4 | 0.5556 | 9 | 5 | 4 | 0.5556 |
| 45 | 16 | 9 | 7 | 0.5625 | 16 | 9 | 7 | 0.5625 |
| >45 |   |   |   | 1.0000 |   |   |   | 1.0000 |

 [↑](#footnote-ref-1)
2. Pooling response data to ensure monotonicity is standard practice when applying the Turnbull estimator and calculating confidence intervals (Haab and McConnell 2002), although excessive pooling can mask issues with the data by understating true confidence intervals (Whitehead 2017). The confidence intervals we report here should be interpreted with caution. [↑](#footnote-ref-2)