NC STATE UNIVERSITY

Center for Environmental and Resource Economic Policy
College of Agriculture and Life Sciences
https://cenrep.ncsu.edu

Costs of Endangered Species Protection on Public Lands: Evidence from Cape Hatteras National Seashore

Steven J. Dundas, Roger H. von Haefen, Carol Mansfield

Center for Environmental and Resource Economic Policy
Working Paper Series: No. 16-016
November 2016

Suggested citation: Dundas, S.J., von Haefen, R.H., and C. Mansfield (2016). Costs of Endangered Species Protection on Public Lands: Evidence from Cape Hatteras National Seashore. (CEnREP Working Paper No. 16-016). Raleigh, NC: Center for Environmental and Resource Economic Policy.

Costs of Endangered Species Protection on Public Lands: Evidence from Cape Hatteras National Seashore

Steven J. Dundas*
Oregon State University

Roger H. von Haefen
North Carolina State University

Carol Mansfield
RTI International

Working Paper Draft: November 2016

Abstract
Management of public lands often involves competing uses and difficult tradeoffs. Here we examine the implications of a direct federal land use conflict in Cape Hatteras National Seashore: off-road vehicle (ORV) access and endangered species protection. Results from a repeated discrete choice model of recreational angler behavior suggest that the economic costs of access restrictions are relatively modest, ranging from $403,000 to $2.07 million annually. Our results provide general support for the National Park Service's recently implemented ORV management plan, as the upper bound of recreation losses is less than a conservative estimate of the benefits of protecting coastal biodiversity.

Keywords: Recreation demand, public land management, endangered species, ORV restrictions, Cape Hatteras

JEL Codes: Q26, Q51

*The authors would like to thank Marty Smith, Laura Taylor, Wally Thurman, Zack Brown, Matthew Godfrey, and Rob Johnston. Additional thanks to participants at the 2016 W-3133 Meeting and the 2015 AAEA Annual Meeting for helpful comments. Carol Mansfield and Roger von Haefen received support from the National Park Service (Order No. T2310081036, GS-10F-0283K). The views expressed in this work belong to the authors and do not reflect the views of the National Park Service.

1. Introduction

Current public land holdings by the United States government span 640 million acres, or more than one quarter of all U.S. land. The Bureau of Land Management (BLM) and the U.S. Forest Service (USFS) manage a majority of these lands for multiple purposes, ranging from natural resource extraction to habitat conservation. There is a long history of conflict between those wishing to utilize public lands for private (e.g. grazing livestock, harvesting timber) and public (e.g. species preservation) benefit. The 2015 National Defense Authorization Act (NDAA) contained over 70 provisions addressing recent public land management decisions, suggesting that conflicting claims on how best to manage public lands can lead to Congressional intervention. Moreover, these conflicts can play out in far uglier ways, as evidenced by the 2016 armed occupation of Malheur National Wildlife Refuge in eastern Oregon.

An increasingly important tradeoff with public land management is balancing recreational access and environmental protection. This tradeoff strikes at the heart of the National Park Service's (NPS) mission to promote both access and protection at the 401 units it manages. A notable recent conflict occurred at Yellowstone National Park, where new NPS rules limiting snowmobile access were adopted in 2013 after a decades-long battle between environmentalists and recreators. Similarly, in the Cape Hatteras National Seashore (CAHA) on North Carolina's Outer Banks, new NPS rules limiting recreational access for off-road vehicles (ORV) were adopted in 2012 to stem negative impacts to endangered species habitat. In this research, we focus on the latter conflict and aim to provide empirical evidence on the likely costs of these ORV rules.

Although ORV use is prohibited on most NPS-managed land, it is permitted in many national seashores where road networks are primitive.[1] In the case of CAHA, there is a long-standing tradition of recreational anglers using beaches as vehicular corridors for accessing the most desirable fishing locations. These beaches also serve as nesting sites for endangered and threatened species protected under existing federal and state law. Despite this, ORV use in CAHA remained largely unregulated until 2008. Since the NPS fully implemented its CAHA ORV management plan in 2012, ORV restrictions have become common, and often limit angler access to the most desirable fishing sites during the most popular fishing seasons. The benefits and costs of these restrictions are at the center of the public lands conflict in CAHA.

Quantifying the non-market tradeoffs of policy interventions on public lands remains a challenging area for economic research. In the context of the aforementioned snowmobile restrictions in Yellowstone National Park, Mansfield et al. (2008) present stated preference evidence that some restrictions are likely to improve overall welfare due to relatively large gains to non-snowmobile users. In the context of ORV trail closures in Colorado, Deisenroth et al. (2009) find relatively modest recreational consumer surplus losses from potential ORV trail closures (under $300,000 per trail per summer) but do not monetize benefits.[2] Englin et al. (2006) and Jakus et al. (2010) also investigate the costs

---

[1] The NPS allows ORV use in only 12 of its 401 units, and 7 of the 12 units are national seashores: Assateague, Cape Cod, Cape Hatteras, Cape Lookout, Fire Island, Gulf Islands, and Padre Islands.
[2] As noted by Jakus et al. (2010), these modest welfare losses are likely over-estimated as the stated preference model used by Deisenroth et al (2009) does not allow for site substitution by recreators if particular trails are closed.

of ORV closures on public lands, yet only Jakus et al. attempt to monetize the welfare costs from trail closures. In their study of Utah federal lands, they find costs on the order of $1.2 million annually.

In this paper, we estimate the effects of access restriction at CAHA beaches on shoreline recreational anglers. Using data from the 2005-2007 Marine Recreational Information Program (MRIP) surveys collected by the National Oceanic and Atmospheric Administration (NOAA), we sequentially estimate participation and site choice models within a repeated discrete choice framework. In our model, individuals residing in the coastal counties of Virginia, North Carolina and South Carolina choose whether, when, and where to engage in coastal recreational fishing. The model is then used to generate upper and lower bound estimates of the behavioral and welfare effects associated with time-varying access restrictions and beach closures implied by the CAHA ORV management plan.

Our results suggest that the annual angler welfare loss due to ORV restrictions range from $403,000 to $2.07 million (2010 dollars), depending on extent of closures in adaptive management areas. The effects on total angler trips in the three-state region are relatively small, with annual visitation rates predicted to decline by 0.6 to 3.6 percent.[3] This finding suggests that most anglers will either continue to take fishing trips to CAHA

[3] Furthermore, a sensitivity analysis shows only a $3.6 million welfare loss annually and a 12.6 percent decline in regional recreational fishing trips if *all* CAHA sites were completely closed.

sites impacted by the ORV restrictions (and thus experience a *diminished* trip) or switch to other non-impacted beaches in the three-state region (and experience a *substitute* trip).

From a policy perspective, our results do not include four factors that could result in larger total costs: 1) increased congestion at beaches that remain open for use; 2) losses from CAHA trips originating from outside coastal counties; 3) enforcement costs; and 4) losses from non-fishing (e.g., surfing) CAHA trips. To address these limitations, we combine our results with previous findings in the literature, CAHA visitor survey results, and anecdotal evidence from the NPS. Using assumptions that imply an upper bound estimate, the losses from CAHA ORV restrictions range from $3.34 million to $12.62 million annually. However, these estimates are still less than conservative lower bound estimates of potential benefits of endangered species protection in coastal North Carolina implied by Whitehead (1993) ($13 million annually) and Dalrymple et al. (2012) ($48 million annually). We therefore conclude that the current ORV restrictions in CAHA likely pass a benefit-cost test, although refinements to current policies could generate higher net benefits.

In additional to providing policy-relevant estimates for an important and timely issue, a notable contribution of our research is to develop new methods to analyze MRIP data. MRIP data is collected with two independent instruments – an on-site intercept survey designed to measure catch and a telephone survey designed to measure participation. Previous research with MRIP data utilizes the intercept survey to model angler trip allocation decisions to coastal counties (e.g. Whitehead and Haab 1999; Whitehead et al. 2009; Hindsley et al. 2011; Alvarez et al. 2014). Here, we use the telephone and intercept survey data to jointly model both participation and trip allocation decisions to individual sites (beaches, piers, marina), thus avoiding potential biases associated with aggregation (Parsons and Needelman 1992). Because not all sites are sampled in every two-month MRIP wave, we develop an algorithm to account for missing sites that leverages auxiliary data used by NOAA when developing its sampling strategy. Finally, our modeling approach is similar to Alvarez et al. (2014) in that we model the angler decision of where (site choice) and when (wave choice) to take a trip. This is necessary to account for spatial and temporal substitution that the CAHA ORV management plan likely generates. Train (2016) has criticized this approach because it introduces unrealistic "time travel," whereby an individual can, for example, shift a summer trip back to spring in response to a summer closure. In our view, the validity of this criticism hinges on whether closures can be anticipated. In contrast to oil spills (Train's point of reference) which generate impacts and closures that are essentially random and unpredictable, CAHA beach closures were announced at least a year before their implementation and thus could be anticipated. Moreover, for those beaches where beach closures were contingent on park managers discovering sea turtle or piping plover nesting sites, NPS announced well in advance that adaptive management plans were in place and could lead to closures.[4] Therefore, when anglers are planning when and where to recreate each year, they can anticipate closures that will be in place at different times and places

[4] The NPS dedicates a portion of their website to FAQ about ORV permitting and closures and utilizes social media (e.g. Facebook) to inform the public on closures in real time, updating anytime a change takes place.

and respond accordingly. This predictability and adaptability mitigates Train's concern in our CAHA beach closure context. Nonetheless, we developed an alternative specification that does not allow for intertemporal substitution and find welfare results that are virtually identical to our preferred specification.

The remainder of the article proceeds as follows. In section II, we present background information on CAHA and the ORV restrictions. Section III outlines our modeling approach and the empirical implementation of the model while Section IV describes our data. Section V discusses results from the site choice and participation models, the policy simulations, and the resulting welfare implications. In section VI, we provide a back-of-the-envelope benefit-cost analysis of the ORV regulations. Section VII concludes.

II. Conflict on Cape Hatteras National Seashore

Cape Hatteras National Seashore was established as the first national seashore in the United States in 1937. The barrier island park stretches over 67 miles, three islands (Hatteras, Ocracoke, and Bodie), and covers 24,470 acres of North Carolina's coastline (see Figure 1).[5] CAHA is located in a relatively remote portion of the Outer Banks, with primary access available from a single bridge (Herbert C. Bonner Bridge on NC Highway 12) on the north end and ferry service to Ocracoke Island on the south end. Since 1989, approximately 2.2 million people have visited CAHA per year. Visitors use the islands

[5] U.S. Fish and Wildlife Service administers 6,000 acres of Hatteras Island as Pea Island National Wildlife Refuge where year-round ORV restrictions are in place.

for a variety of recreation activities, including shoreline fishing. Recreation visits generate a robust local tourism industry that supports eight unincorporated villages on the islands (Landry et al. 2016). Access to many of the park's prime locations for recreation activities often requires ORV beach driving due to primitive road networks. ORV users must purchase a permit ($120 and $50 for annual and seven-day permits, respectively), access the beach at designated dune crossovers, and drive seaward of the primary dunes (i.e., driving on dunes is prohibited).

Concomitant with growing recreational visitation and ORV use in recent decades, CAHA also experienced a decline in the population of nesting shorebirds, including the piping plover (*Charadrius melodus*). This plover is a protected species under the federal Endangered Species Act and is designated as threatened.[6] Piping plovers are highly vulnerable to habitat loss from increasing shoreline development and nesting disruption from human activities. The transitional over-wash areas along the shoreline are prime shallow-sand nesting sites for these plovers and directly overlap with many ORV routes. CAHA also provides critical nesting habitat for two species of sea turtle: loggerhead (*Caretta caretta*), and green (*Chelonia mydas*).[7] Loggerhead turtles visiting North Carolina beaches are classified as both endangered and threatened depending on the sub-population of the turtle, while the green turtle populations retain a threatened status. For

---

[6] The shorebird's population along the Atlantic Coast of the U.S. and Canada was estimated in 2011 at 1,762 nesting pairs (USFW 2011).

[7] Leatherback (*Dermochelys coriacea*), hawksbill (*Eretmochelys imbricata*), and Kemp's ridley (*Lepidochelys kempii*) sea turtles are also found in the waters off North Carolina but nesting on CAHA is not common.

both species, North Carolina represents the northern limits of their nesting grounds.[8] Nesting activity is vulnerable to coastal armoring (e.g. jetties and sea walls), nighttime activity on the beach, and ORV use in nesting areas. Monitoring sea turtle nesting sites began in 1987 when only 11 were found.

The NPS is tasked with managing both recreational access and species protection in CAHA. Their adaptive management plans are under the purview of Executive Orders (E.O. 11644 of 1972 and E.O. 11989 of 1977) and federal laws, including the Endangered Species Act, The Migratory Bird Treaty Act, and the National Environmental Policy Act. On the recreation side, ORV culture is deeply ingrained with local residents and visitors to the area. This stems from the historical use of beaches as a transportation conduit and the lack of ORV access or use restrictions until recently. The Interim Protected Species Management Strategy was drafted in 2007 to manage ORV use but was met with opposition from both environmental groups and ORV advocates. The NPS was sued by wildlife advocacy groups (The Audubon Society and Defenders of Wildlife) that claimed the interim rules did not do enough to protect nesting sites. A settlement was reached in April 2008 that allowed implementation of a temporary strategy including large buffers around sensitive areas and restrictions on nighttime ORV use during the sea turtle nesting season.[9]

---

[8] This is biologically important, as cooler sands produce more male hatchlings, making North Carolina a critical breeding ground for the male populations of each species.
[9] This restriction appears to be helping as NPS has found an average nesting total in CAHA around 129 annually from 2008 – 2011.

The final management rule went into effect on February 15, 2012. The NPS chose between a number of alternative plans after a period of public comment (see National Park Service 2010; Mansfield et al. 2011). The implemented plan was *Alternative F*, which provides a balanced approach between ORV access and vehicle-free areas relative to other alternatives considered. Of the 67 miles of coastline in CAHA, *Alternative F* designates 27.9 miles for year-round ORV routes, 12.7 miles of seasonally accessible routes, and 26.4 vehicle-free miles. Included with this alternative are planned infrastructure improvements including parking lots at key locations and improvements to a sand road system outside of nesting areas. Wildlife management areas and village beaches (areas directly adjacent to population centers) are closed to allow for shorebird breeding activity, typically March to July. This alternative re-opens these areas earlier and for longer periods of time than the other alternatives considered by NPS. Nighttime restrictions on all beaches are in place May 15th to September 15th from one hour after sunset until cleared by patrol in the morning. The map in Figure 2 illustrates the spatial variation in restrictions in place on June 1st, 2015, on Hatteras Island. The environmentally preferred alternative (*Alternative D*) was also considered and this plan limited ORV use to designated year-round routes with minimal new construction of ramps and no new parking areas. In comparison to *F*, *Alternative D* would designate 27.2 miles for year-round ORV routes, zero miles of seasonally accessible routes, and 40.8 vehicle-free miles. Nighttime ORV use would be prohibited between 7 PM and 7 AM from May 1st to November 15th to protect nesting turtles. Permit fees would be lower under this plan due to the decreased need for NPS oversight and management.

It is important to note that the CAHA final ORV rule is again being challenged due to a late addition rider to the 2015 NDAA. Over 70 public land provisions were added to this omnibus legislation, including language forcing a review of the CAHA ORV restrictions. The Secretary of Interior is now required to investigate the potential for reducing the size of nesting buffers, opening beaches earlier in the morning during the summer, extending ORV routes and access points, and modifying the size and location of restricted areas in CAHA. In June 2015, the NPS proposed changes to existing ORV restrictions that would involve: 1) conducting pre-dawn beach patrols to promote earlier beach openings; 2) expanding the ORV driving season by four weeks (two in the spring, two in the fall), and 3) limiting vehicle free areas and allow more seasonal use. After five public hearings in August 2015 and further internal deliberations, NPS published their proposed revision to the CAHA management plan in the Federal Register in August 2016. A final decision is likely in early 2017. Although we consider only the existing ORV rules in this research, the proposed revisions will, if anything, reduce the regulatory costs and strengthens our overall finding that the net-benefits of the ORV restrictions are positive.

III. Modeling Strategy

A distinctive characteristic of CAHA ORV restrictions is that they vary across space and time. In other words, sections of CAHA may be closed to ORVs in certain months and open in others. When monetizing the non-market losses from these policies, accounting for two types of behavioral responses seems particularly important: 1) spatial and temporal substitution of trips across recreational sites and months of the year; and 2) choosing to not take a trip (i.e., participate) in response to a closure. To allow for both behavioral responses, we use the repeated discrete choice, random utility maximization (RUM) framework (Morey et al. 1993) to model whether, where, and *when* anglers engage in shoreline recreational fishing. An individual first chooses whether to participate in coastal fishing and then conditionally chooses which site to recreate at as well as which two-month period (or wave) to go. This structure, which is illustrated in Figure 3, is similar to Alvarez et al. (2014) who consider beach closures resulting from the 2010 Deepwater Horizon Oil Spill but differs from most discrete choice applications that abstract from the timing of a trip within a recreation season. In our application, modeling when recreational trips occur is important because site-specific ORV restrictions vary across the recreation season and a plausible response to a closure at a particular beach is to substitute the trip to a different time of year when the same beach is open to ORV use. Since the MRIP data we employ temporally disaggregates trips to a site by wave (January-February, March-April and so forth), we treat the objects of choice in our model as all site/wave pairs and a "no trip" alternative. As discussed in the introduction, we believe this specification of the choice set is not subject to Train's (2016) "time travel" critique, but for robustness, we also consider a specification that only allows for spatial substitution within each two-month period. Sites and the "no trip" alternative are the objects of choice in this case.

In the RUM framework, an individual chooses the alternative that maximizes her utility. The factors that drive individual decisions can be separated into determinants that are either observable or unobservable, with the latter treatable as random from a modeling perspective. The conditional indirect utility of individual *i* from choosing site *j* and wave *w* on choice occasion *t* is specified as follows:

$$V_{ijwt} = U(m_{it} - c_{ijw}, \mathbf{X}_{jwt}, \varepsilon_{ijwt}) \tag{1}$$

where m_{it} is income, c_{ijw} is travel cost,[10] $\mathbf{X}_{jwt}$ is a vector of site characteristics, and ε_{ijwt} captures idiosyncratic, random factors. Conditional on taking a trip, a rational angler selects the site *j* and wave *w* that generates the highest utility. More precisely, angler *i* chooses site *j* in wave *w* for their recreation activity if $V_{ijwt} \geq V_{ikw't}, \ \forall \ k, w'$. For convenience, we assume utility is linear and additive in ε_{ijwt} (i.e., $V_{ijwt} = v_{ijw} + \varepsilon_{ijwt}$).

As discussed in a later section, the MRIP data is collected with two separate and independent surveys that provide repeated cross-sectional information on site choice and participation. The survey sampling protocols imply that we do not observe both where and how often a given individual recreates, and thus we are restricted to econometric models that completely separate these two dimensions of choice. Therefore, we employ a two-level nested logit model (Morey 1999) and sequentially estimate a conditional

[10] Because our income and per mile driving cost data only varies annually, we make the further assumption that travel costs are equal across waves within a year for each individual/site pair.

site/wave choice model and a participation model. The model assumes the errors are independently drawn from a generalized extreme value distribution, which essentially allows a common random component to enter the site-specific errors. This induces correlations in the conditional indirect utilities for each site/wave pair and more reasonable substitution patterns. By sequentially estimating site choice and participation decisions, we can recover a complete characterization of recreation behavior.[11]

The probability of choosing site *j* and wave *w* on choice occasion *t* is:

$$P_{ijwt} = P_{it}(j, w \mid trip) \times P_{it}(trip)$$
$$= \frac{e^{(v_{ijw}/\lambda)}}{\sum_{w=1}^{W}\sum_{j=1}^{J} e^{v_{ijw}/\lambda}} \frac{\left[\sum_{w=1}^{W}\sum_{j=1}^{J} e^{v_{ijw}/\lambda}\right]^{\lambda}}{e^{v_{i0}} + \left[\sum_{w=1}^{W}\sum_{j=1}^{J} e^{v_{ijw}/\lambda}\right]^{\lambda}} = \frac{e^{(v_{ijw}/\lambda)}\left[\sum_{w=1}^{W}\sum_{j=1}^{J} e^{v_{ijw}/\lambda}\right]^{\lambda-1}}{e^{v_{i0}} + \left[\sum_{w=1}^{W}\sum_{j=1}^{J} e^{v_{ijw}/\lambda}\right]^{\lambda}} \qquad (2)$$

where λ is the dissimilarity coefficient, bounded by theory between 0 and 1, and *W* and *J* are the total number of waves and sites. If we assume the conditional indirect utility of not taking a trip (alternative 0) is $V_{i0t} = v_{i0} + \varepsilon_{i0t}$ with ε_{i0t} an independent draw from the Type I extreme value distribution, the probability of not taking a trip is then:

[11] The large datasets used here suggest that efficiency loss relative to full-information maximum likelihood estimation is relatively small. Furthermore, the data do not allow estimation of a random coefficient or latent class model due to the calibration step necessary for the alternative specific constants to model each site individually.

$$P_{i0wt} = \frac{e^{v_{i0}}}{e^{v_{i0}} + \left[\sum_{w=1}^{W} \sum_{j=1}^{J} e^{v_{ijw}/\lambda} \right]^{\lambda}}. \tag{3}$$

All parameters of this model can be recovered by first estimating the site/wave choice model and then conditionally estimating the participation model using standard logit estimation techniques (Ben-Akiva and Lerman 1985).

Empirical Implementation of the Model

Estimation of model parameters proceeds in three steps. Step 1 utilizes the MRIP intercept data to estimate a conditional logit site/wave choice model with a full set of alternative specific constants (ASCs) separately by year. The conditional indirect utility (i.e., equation 1) from individual *i* visiting site *j* in wave *w* on choice occasion *t* is defined as follows:

$$V_{ijwt} = \beta c_{ijw} + \delta_{jw} + \varepsilon_{ijwt} \tag{4}$$

where we employ the common assumption of a constant marginal utility of income. The site choice and wave probabilities can then be written:

$$P_{it}(j, w \mid trip) = \frac{e^{\left(\beta c_{ijw} + \delta_{jw}\right)/\lambda}}{\sum_{w=1}^{W} \sum_{k=1}^{J} e^{\left(\beta c_{ikw} + \delta_{kw}\right)/\lambda}} \tag{5}$$

where δ_{jw} is an ASC specific to each site/wave pair. The ASCs are designed to capture site and wave specific characteristics that vary across years (e.g., catch rates, species composition) but are common across sites during a given wave. Although this approach does not explicitly include site characteristics, the ASCs will nonparametrically control for all characteristics that do not vary across individuals within each wave. Note the first stage estimation does not permit separate identification of the dissimilarity coefficient (λ) from (β, δ_{jw}). Rather, the ratios β / λ and δ_{jw} / λ are estimated in step 1. As described below, we lay out a strategy for estimating λ separately in step 3, which in turn allows us to back out β and δ_{jw}.

This first step generates consistent estimates for the normalized travel cost coefficients (β / λ). However, since the MRIP survey samples at most 40 percent of fishing sites in each wave/year, this step does not generate consistent estimates for the ASCs because ASCs are not identified for unsampled sites. Therefore, our second step uses auxiliary data on aggregate trip frequency at every site to calibrate all ASCs. The auxiliary data comes from site registry files that contain "fishing pressure" estimates which NOAA utilizes when designing its sampling protocols. The fishing pressure estimates vary across sites, waves and years and correspond to NOAA's best estimate of the number of site visitors in a normal 8-hour period at the time sampling commences.[12]

[12] NOAA continuously updates these fishing pressure estimates based on feedback from infield staff, news reports and other sources. Admittedly the estimates are ex ante measures of fishing intensity and subject to

These estimates reflect visitation rates for individuals living inside and outside coastal counties. To generate visitation estimates for only those individuals residing in coastal counties which, due to coverage limitations with the phone survey, are the focus of our analysis, we used the intercept data to construct estimates of the share of trips originating from coastal counties to adjust the site registry estimates. These adjusted visitation estimates are then converted to trip shares and Berry's (1994) contraction mapping is used to recover calibrated estimates of the ASCs (see Appendix A for further discussion of the site registry data and calibration procedure). The adjusted ASCs are then used to generate the inclusive value index:

$$IV_i = \ln\left(\sum_{w=1}^{W}\sum_{k=1}^{J} e^{\left(\beta/\lambda c_{ikw} + \delta_{kw}\right)}\right) \tag{6}$$

where β/λ is estimated in step 1 and $\tilde{\delta}$ is estimated in step 2. The inclusive value term can loosely be interpreted as the expected utility of a trip (Hausman et al. 1995).

The third and final step estimates a standard discrete choice logit model of participation as a function of the inclusive value, demographics, a 300-mile proximity to CAHA indicator,[13] and area code and year fixed effects. We assume the total number of

forecasting errors. Nonetheless, they have complete coverage for every site, wave and year and represent the best estimates available for shoreline recreational fishing activity in the region.

[13] This dummy variable is designed to control for the possibility that individuals living in the three-state area but relatively far away from CAHA are likely to have recreation sites that are not in our analysis. For example, residents of Northern Virginia are likely to consider sites in Maryland and Delaware when making coastal fishing trips, but these sites are not represented in our model. To control for this incomplete coverage of relevant sites, we construct this dummy variable which identifies individuals residing more than a 300 mile one-way driving distance from CAHA.

choice occasions on which an individual may take a trip is proportional to the length of the reporting period scaled by an adjustment factor that ensures all respondents have more choice occasions than trips. The utility function associated with non-participation is defined as:

$$V_{i0t} = \delta_0 + IV_i\lambda + P_i\omega + \mathbf{Y}\alpha + \mathbf{A}\tau + \boldsymbol{D}_i\phi + \varepsilon_{i0t} \tag{7}$$

where $\boldsymbol{Y}$, $\boldsymbol{A}$ and $\boldsymbol{D}_i$ are vectors of year dummies, area code dummies, and demographics, respectively, and P_i is the CAHA proximity indicator equal to 1 if the individual is outside a 300 mile one-way driving distance of CAHA fishing sites. The recovered parameter estimate on IV_i is the dissimilarity coefficient (λ), which can be multiplied by the first-stage normalized travel cost coefficients to recover a consistent estimate of β.

IV. Data

MRIP data from NOAA’s National Marine Fisheries Service (NMFS) used for this analysis were collected with the point-of-access Angler Intercept Survey and the Coastal Household Telephone Survey. Table 1 provides a concise summary of all variables used in each stage of the analysis. For the intercept data, observations are restricted to individuals fishing at shoreline sites in Virginia, North Carolina, and South Carolina from 2005-2007. This timeframe is representative of the region *before* new ORV restrictions were implemented. The data are compiled in two-month intervals, resulting in six waves per year. This survey collects data from interviewed individuals on their catch and mode of fishing. The primary variable of interest for this work is the zip code of residence for each survey respondent. This information allows estimation of travel costs from their home to the fishing site where they were intercepted as well as other sites in their choice set. Note that each individual is interviewed only once so the resulting dataset is a repeated cross-section. Unlike previous studies that aggregate sites into counties (e.g., Alvarez et al. 2014), we include all 344 individual MRIP sites along the Atlantic coasts of Virginia, North Carolina, and South Carolina in our analysis. Across the 18 waves spanning 2005 to 2007, we have data for 17,559 shoreline fishing trips originating from 4,657 different zip codes that are included in the conditional site/wave choice model in step 1 described in the previous section.

Consistent estimation of our nested logit model requires an adjustment for the stratified sampling design. The intercept survey is stratified by site, state, mode, year and wave, implying that where anglers fish is correlated with their likelihood of inclusion in the sample. In 2012, NMFS published design-based sampling weights for all intercepted trips dating back to 2004. By construction, these weights produce unbiased estimates of angler effort and reflect the proper proportion of trips from coastal and non-coastal origins (Breidt et al. 2012; Lovell and Carter 2014). Use of these weights obviates the need for econometric solutions to endogenous stratification (e.g. Hindsley et al. 2011).

Calculation of the travel costs for each observation involves several steps. First, *PC*Miler* calculates the one-way driving distance (*dist*), travel time (*time*), and tolls (*toll*) from an origin zip code centroid to all potential shoreline fishing sites in a given choice set. Next, additional data are obtained on average fleet fuel economy (*fe*) from the U.S. Department of Transportation, gas prices by state (*gas*) from the US Energy Information

Administration, automobile per-mile operation costs (*cpm*) from AAA, and zip code level income from the U.S. Census Bureau. To construct travel costs for each individual survey respondent, costs that can be shared by all persons on a given trip (e.g. tolls, gas, and mileage) are divided by the average number of individuals in each party from our sample ($\bar{n}$).[14] The opportunity cost of time (*oppc*) is determined using the common assumption of 1/3 of the wage rate (Cesario 1976). Round trip travel costs (in 2010 dollars) for individual *i* to site *j* on choice occasion *t* are calculated as:

$$c_{ijwt} = 2*\left(\begin{bmatrix} dist_{ij}*(cpm_t)+ \\ (dist_{ij} / fe_t)*gas_t \\ +toll_{ij} \end{bmatrix} / \bar{n} + oppc_{it}*time_{ij}\right) \tag{8}$$

From the phone survey, information is collected by county-stratified random-digit-dial (RDD) from coastal households from 2005 to 2007 on the frequency of fishing trips in the preceding two months.[15] The data include the number of anglers who have taken trips and the number of trips taken by each angler in the previous two months. For this study, the phone survey pulls from coastal counties in the three-state region surrounding the Outer Banks of North Carolina with six-digit phone exchanges as the spatial unit of analysis. Data from 4,928 phone exchanges are included. Survey records for individuals who actively participated in coastal recreation fishing are used to characterize participation in the model. However, the non-participating households are only identified

---

[14] The intercept data contains a variable for the number of people per fishing party. The average is 2.73.
[15] The 2005-2007 timespan is also ideal to limit potential impacts on representativeness of RDD surveys due to growing cell phone use and recent decreased willingness to participate in phone surveys.

at the county level – a less refined spatial scale then the fishing households. Since the phone survey was RDD within each county, we disaggregate the county level non-fishing sample to the phone exchange level to facilitate analysis in the following way. Each relevant six-digit phone exchange is assigned a population weighted proportion of the count of non-fishing individuals in the county where the exchange in located. For example, assume a county with three phone exchanges, each with a population of 1,000 people. If the RDD survey contacted 30 non-anglers in the county, randomization implies we can assign 10 non-anglers to each phone exchange in that given two-month period.

As noted earlier, the third step of model estimation includes demographics, year and area code dummies and a proximity indicator as covariates affecting the participation decision. Zip code level demographic data on average household income, race, sex, population density, and education are gathered from U.S. Census Bureau's American Community Survey. Zip codes are linked to six-digit phone exchanges using a proprietary data set from *Melissa Data*.[16] A population-weighted average of zip code demographic data is assigned to each phone exchange.

V. Results

Estimation of the conditional site/wave choice model yields travel cost coefficients that are highly significant and negative as expected across all years of analysis (see panel A of

[16] Dataset description available here: http://www.melissadata.com/reference-data/fonedata.htm

Table 2). Four sets of results for the logit participation model are reported in panel B of Table 2. The first column reports results from estimating the participation decision with demographic data and a single inclusive value (Model 1). Results from Model 2 with year-specific inclusive values are very similar to the previous specification so we return to the more parsimonious single inclusive value in Models 3 and 4.[17] Model 3 adds the CAHA proximity indicator and then the preferred specification (Model 4) further adds year and area code fixed effects to control for time trends and any location–specific time-invariant unobservables that may impact the participation decision, respectively.

Results from our preferred specification suggest that participation is likely to increase with income and decrease when anglers reside more than 300 miles from CAHA. Coefficients on education, gender, and population density are no longer significant in this specification likely due to the addition of the area code fixed effect. The estimate for the dissimilarity coefficient is 0.04, which falls within the 0-1 interval and thus is consistent with RUM theory (Herriges and Kling 1997). This result implies a high degree of correlation among the site choices in each nest and is relatively low when compared similar estimates in the literature at large. The implication here is a relatively large per trip value of approximately $342 as compared to results from two recent meta-analyses

---

[17] Largely for computational reasons, we do not correct the second-stage standard errors to account for the fact that generated inclusive values are constructed with travel cost parameters that are econometrically estimated in the first stage. Because these parameters are tightly estimated, however, we do not suspect that doing so would imply substantially different inference.

that included saltwater fishing trips.[18] Moeltner and Rosenberger (2014) report the average WTP/day for a saltwater fishing trip in the Northeast is $39.39 (2010 dollars) from five relevant valuation studies. Johnston and Moeltner (2014) show mean WTP/day from 14 different studies for saltwater fishing of big-game species is approximately $33.06. They also report the average WTP/day for small-game saltwater fishing across 13 studies as $21.33.

Given our estimate for the dissimilarity coefficient implies a much larger value for a fishing trip than previous work, we identify a potential source of measurement error that may be driving this result. We suspect that the imprecise nature of the trip origin information in the phone survey data (i.e., respondents spatially identified by phone exchange, not zip code) is introducing measurement error into the inclusive values which in turn likely generates attenuation bias with the estimated dissimilarity coefficient. Moreover, the fact that the ASCs that feed into the inclusive values are calibrated with fishing pressure data in the site registry, and not precisely estimated with choice data, may introduce additional measurement error.

In order to test the implications of this potential bias, we run policy simulations with our estimated model and with a model that calibrates the dissimilarity coefficient using a $30 per trip value, which is consistent with the average empirical meta-analyses reported

[18] This estimate is constructed using the formula $-1/\beta$ for a value of trip, where β is estimated by taking the product of the mean value of the first-stage normalized travel cost estimate (-0.073) and the estimated dissimilarity coefficient (0.04). See Haab and McConnell (2002) for a derivation of this result. This approximation should be relatively robust given our large choice set application.

in Moeltner and Rosenberger (2014) and Johnston and Moeltner (2014). As shown by Haab and McConnell (2002), the value of a fishing trip in a repeated RUM framework is approximately -1/β, where β is the travel cost coefficient. From our first stage trip allocation model, we have an estimate of β/λ. Therefore, if we impose a \$30 value of a trip, we can infer the calibrated value of λ (0.46). As we describe below, our policy implications are robust to either approach.

Policy Simulations

We now investigate the behavioral and welfare effects on shoreline recreational anglers associated with two alternative ORV access management plans. As described in a previous section, *Alternative F* is the NPS preferred alternative that was eventually implemented in 2012 and *Alternative D* is the environmentally preferred alternative. There are 16 MRIP survey sites in CAHA that would be impacted by either strategy. Table 3 displays the different wave-specific management strategies at these sites under *Alternative F,* and Table 4 summarizes the assumptions used to model the closures.[19] A letter 'O' indicates that the site remains open to all activities, including ORVs. 'X' indicates that a site is closed and both ORV and pedestrian use is prohibited. 'XP' indicates that ORVs are prohibited but the site is still open to pedestrians. Lastly, 'A'

[19] The management plan and closure scenarios for *Alternative D* are provided in Table B.1 in the Appendix.

indicates adaptive planning areas where a site could either be open (‘O’) or be closed for resource protection (i.e. ‘X’ or ‘XP’).

The policy scenarios for site/wave choice combinations ‘O’ and ‘X’ are straightforward to model – sites remain open or are closed. Modeling ‘A’ and ‘XP’ is more challenging and thus necessitates estimating a range of possibilities (i.e., an upper and lower bound) dependent on the specifics of the management strategies. For ‘XP’ sites, pedestrian access is allowed, but individuals would have to walk considerable distance carrying their gear to reach the fishing site (0.5 – 2 miles one way) from the closest available parking lot. In some cases, the distance to cover on foot would be excessive, so the site would essentially be closed (i.e., modeled as ‘X’) and represent the upper bound of potential costs. Alternatively, if a site is reasonably accessible by foot, two hours of round-trip travel time are added to account for additional opportunity cost of time needed to access the site. This represents the lower bound of potential costs under ‘XP’ restrictions. The range of options for the adaptive strategy ‘A’ is more direct. The upper bound is modeled as if the site is closed (‘X’) and the lower bound as if the site is open (‘O’).

As noted earlier, nighttime driving restrictions were first imposed in 2008 so the primary policy simulation analysis uses our model estimates with data from 2005 – 2007 as the baseline before any ORV restrictions were imposed. ASCs and inclusive values (equation 5) are estimated under each alternative management plan at both the upper and lower bound. The change in WTP for shoreline recreational fishing for individual *i* under each policy scenario is estimated using the following equation (Haab and McConnell 2002):

$$WTP_i = -\frac{T_i}{\beta}\left(\ln\left(e^{-v_{i0}} + \left[\exp(IV_i^1)\right]^{\lambda}\right) - \ln\left(e^{-v_{i0}} + \left[\exp(IV_i^0)\right]^{\lambda}\right)\right) \qquad (9)$$

where β is the travel cost parameter, T_i is the number of choice occasions, IV_i^0 is the inclusive value for individual *i* in the baseline period and IV_i^1 is inclusive value under the policy scenario. In equation (9), the differences in inclusive value terms from the baseline to each policy scenario drive the differences in WTP.

Our preferred policy simulation uses our site/wave choice model allowing both spatial and intertemporal substitution with the calibrated dissimilarity coefficient. Welfare changes are reported in Table 5 and the demand responses in Table 6. Standard errors for all estimates are generated with a parametric bootstrap (Krinsky and Robb 1986) with 100 draws taken from the estimated parameter vector and covariance matrix. All predictions for welfare and demand responses are highly significant, as suggested by the 95% confidence intervals reported in the tables.

The primary result is the relatively modest loss estimates predicted under both management alternatives. The total range of welfare losses (in 2010 dollars) is $403,000 (*Alternative F*, lower bound) to $2.75 million (*Alternative D*, upper bound) annually. The annual welfare loss under the most restrictive scenario (upper bound) of the current management plan (*Alternative F*) is estimated to be approximately $2.07 million per year. Second, the projected incremental cost associated with moving from the current plan to the environmentally preferred *Alternative D* (i.e., more protections in place for threatened wildlife**)** ranges from $294,000 to $680,000 annually.

Third, long-run demand responses to the policies predicted in the simulations are also relatively modest. These results support the notion that anglers will adapt to the ORV restrictions and choose alternative sites for shoreline recreational fishing activities. It is important to note that a site's closure for resource protection does not imply a complete loss of recreation. Individuals have the ability to substitute to other locations unaffected by the management policy. Alternatively, they can continue to take a diminished trip to an affected site (i.e., closed to ORVs but still open to pedestrian traffic) if they are willing to bear additional costs associated with accessing the site by foot. For each year, the model is utilized to estimate of the number of *affected* trips (i.e., trips impacted by the ORV restrictions), *lost* trips (trips that occur in the baseline scenario but not with ORV restrictions in place), and *diminished* trips (trips to impacted sites that continue to occur with ORV restrictions but generate less utility to anglers) as a result of the ORV restrictions. *Substitute* trips, or trips that were shifted to non-impacted sites after the ORV rules were implemented, are then calculated using the following identity:

$$\textit{Affected Trips} = \textit{Lost Trips} + \textit{Diminished Trips} + \textit{Substitute Trips} \qquad (10)$$

Approximately 143,000 (7.5 percent of the total 1.9 million) trips per year are affected under *Alternative F*. The upper bound on the number of lost trips is 3.6 percent of total trips in the three-state region, or approximately 68,000 trips. Switching from *F* to *D* at the upper bound could increase the decline to 6.2 percent, or an additional loss of 50,000 trips. Of the total affected trips, 75 to 81 percent are diminished trips in the lower bound simulations. By construction, there are zero diminished trips at the upper bound as all sites that have the potential to be closed are closed. At the lower bound, substitute trips represent approximately 10 to 12 percent of affected trips, with the aggregate numbers being relatively similar to the numbers of lost trips. When diminished trips are not possible in the upper bound scenarios, substitute trips account for slightly more than half of the affected trips. If adaptive management areas remain accessible to pedestrians (lower bound), diminished trips account for a majority of the affected trips with lost trips being less than 9 percent and 12 percent of the total for *Alternatives F* and *D*, respectively.

Given the “time travel” critique of Train (2016) of our preferred site/wave choice model and our decision to calibrate the dissimilarity coefficient, we estimate a number of model variations, including an uncalibrated site/wave choice model and both calibrated and uncalibrated site choice only models (i.e. no intertemporal substitution), as sensitivity checks on our main results. The welfare results from these models under *Alternative F* are displayed in Table 7. The losses are remarkably similar with all welfare point estimates failing within the 95% confidence interval of our preferred model. The upper (lower) bound estimates from the sensitivity checks are within 0.1 – 5.3 (0.7 – 10.7) percent of the preferred model estimates. These results support our model choice that captures both spatial and temporal substitution behavior and our calibration decision to account for measurement error in our survey data. Full reporting of the simulation results using the alternative model specifications are provided in the appendix.

Potential for Additional Costs

This section identifies four potential limitations to our main welfare results that may imply additional costs associated with CAHA ORV restrictions: 1) increased congestion at beaches that remain open for use; 2) losses from CAHA trips originating from outside coastal counties; 3) losses from non-fishing (e.g., surfing) CAHA trips; and 4) enforcement costs. Below we discuss these limitations and our subsequent assumptions to calculate an upper bound, or "worst-case" scenario, of the potential costs that incorporate them.

Given that ORV use is prohibited at some CAHA fishing sites and anglers are potentially substituting to other sites, increases in congestion at sites that remain open is a salient concern. Welfare estimates presented above may be biased downward as they ignore any lost angler utility from recreating at more congested sites. Recent empirical work has highlighted the issue of potentially understating costs of closures due to congestion. Timmins and Murdoch (2007) address endogeneity concerns related to congestion with an instrumental variables approach and their results suggests that congestion may increase the costs of closures by 50 percent or more. Bujosa et al. (2015) take a site density approach to improve the measurement of congestion and also find increases in costs of approximately 50 percent when site congestion levels are adjusted after closures. Therefore, we add a 50 percent increase in our cost estimates to account for the potential adverse effects on utility of congestion at sites that remain open during ORV closures.

Second, the lost welfare estimates from our econometric model only apply to local anglers residing in coastal counties as defined by the MRIP survey. Over the time period studied (2005-2007), MRIP intercept data indicate that anglers who live in these areas are responsible for approximately 54 percent of the user days to sites in the Outer Banks of North Carolina. This implies that 46 percent of the user days are from anglers residing in non-coastal counties that are not represented in the above welfare analysis. Comparison of coastal and non-coastal angler user days in the weighted MRIP intercept data reveal that number of hours fished, catch rates, number of fish landed, and mode of fishing is relatively similar between these groups at varying spatial scales – three-state region, North Carolina, and Dare County, NC (see Table 8). Based on these findings, we conclude that it is plausible to assume an equal value of user days for both local and non-local anglers. Therefore, we scale up total welfare losses to account for non-coastal shoreline recreational anglers.

Third, shoreline fishing is not the only recreation activity pursued by ORV users. A NPS-sponsored user survey on CAHA conducted in 2009-2010 (Mansfield et al. 2010) found that shoreline fishing represents a large share of recreation activities impacted by the ORV rule, as 38 percent of respondents reported "beach fishing" as an activity they had taken part in during their current trip. "Swimming, sun-bathing, or enjoying the beach," "bird-watching," and "surfing" are other recreation activities garnering relatively high percentages in the survey. However, for our purposes, this estimate is of limited value because respondents include both ORV and non-ORV users of CAHA. Nonetheless, our sense is that anglers remain the primary recreation group impacted by the ORV restrictions as evidenced by: 1) the avid opposition to the rules by fishing clubs (Williams 2012); and 2) anglers representing the large majority of individuals present to voice opinions at recent NPS public meetings to discuss modifications to the current ORV rules. Additionally, anecdotal evidence from our discussions with NPS staff, ORV users, and journalists suggest that a plausible assumption for an upper bound on the percentage of ORV users that are *not* anglers is about 50 percent. Therefore, we then double the welfare effects to account for potential non-angler ORV-based recreation losses.

Lastly, the new rules on CAHA for ORV use require purchasing a permit to drive on the beach. An annual pass can be purchased for $120 and a seven-day pass for $50. Since 2012, the Park Services has sold approximately 10,000 annual passes and 20,000 daily passes each year. This represents an additional cost to recreators of approximately $2.2 million. These fees are used to cover enforcement costs within CAHA but also to maintain and build parking lots and access points for ORV users. For example, the "Inside Road", a 4 mile long interdunal sand road designed to improve access to popular fishing areas, was completed from Buxton to Frisco in January 2016 using funds from these permits. For our purposes, we assume the net benefits of this infrastructure is zero and that half of the fees (i.e., $1.1 million) are social costs associated with enforcement of the ORV management plan.

To illustrate the potential magnitude of additional costs stemming from these four assumptions, let us consider a simple exercise starting with our upper and lower bound cost estimates under *Alternative F*. Our model estimates a welfare loss to local anglers of approximately $403,000 to $2.07 million per year. The inclusion of congestion costs at sites remaining open has the potential to increase the range of welfare losses to $605,000 and $3.11 million. The addition of non-local anglers could raise the potential scope of welfare losses with congestion to $1.12 million - $5.76 million annually. Then, if this number is scaled up to include all other ORV-based recreation, total welfare losses have the potential to reach $2.24 million to $11.52 million. Adding in the enforcement costs brings the total range to $3.34 - $12.62 million annually. This back-of-the-envelope exercise is summarized in Table 9. It is important to note that this range is likely overstating the potential increases in welfare loss compared to the model estimates given the upper bound or "worst-case" scenario implications of our assumptions.

VI. DISCUSSION

To put the welfare loss estimates from our analysis and the back-of-the-envelope calculation above into perspective, it is important to consider the benefits for protecting coastal biodiversity (i.e., piping plovers and sea turtles) that are motivating the ORV restrictions. Economic estimates for such non-market goods are typically generated with stated preference survey methods and the most relevant for the analysis here is Whitehead (1993) and Dalyrmple et al. (2012). The Whitehead study presents results from a contingent valuation survey of wildlife preservation programs in coastal North Carolina. The research estimates an annualized household willingness to pay (WTP) for loggerhead sea turtle preservation of $10.98 and for all non-game endangered and threatened species in coastal North Carolina (including piping plovers and loggerhead sea turtles, among others) of $14.74 (both in 1991 dollars).[20] Dalrymple et al. (2012) find household WTP for non-game endangered species (i.e., piping plovers, sea turtles, among others) ranging from $65 to $98.80 per year. Aggregating these estimates to all 3.7 million North Carolina households and adjusting them into 2010 dollars produces aggregate benefits from Whitehead of $65 million to $87.3 million annually and $240 million to $366 million annually from Dalrymple et al. According to N.C. Wildlife Resources Commission's seaturtle.org (September 30, 2015) website, approximately 20 percent of North Carolina loggerhead nests occur in CAHA. Multiplying the *lower* bound aggregate benefit estimates from each study by one-fifth to bring the spatial scale of the benefits in line with the costs yields annual benefits of $13 million (Whitehead) and $48 million (Dalrymple et al.). Thus, the long-run benefit cost ratio of the ORV restrictions ranges from 1.03 to 3.8 even when comparing the maximum of the upper bound of potential costs ($12.62 million) to conservative, lower bound estimates of the benefits.

It should be recognized that our analysis does not include *all* the costs and benefits associated with ORV restrictions in CAHA. First, we do not include lost profits to local businesses (e.g., shops, hotels and restaurants) from a reduction in visitors. Given the vocal opposition to ORV restrictions by local businessmen, these losses very well may be

[20] The contingent valuation questions were asked as follows: "Suppose that a $A contribution from each North Carolina household each year would be needed to support and fund the loggerhead sea turtle program (nongame wildlife management program). Would you be willing to contribute $A each year to the 'Loggerhead Sea Turtle Preservation Trust Fund' ('Coastal Nongame Wildlife Preservation Trust Fund') in order to support the loggerhead sea turtle program (nongame wildlife management program)?" where A = {1,5,10,25,50,100}.

substantial. But from a strict benefit-cost perspective, any lost profits associated with diminished economic activity in CAHA should be offset by increased profits at non-CAHA businesses to the degree that visitors respond to the ORV restrictions by spending their dollars elsewhere. Ultimately, the overall effect on economic activity is an empirical question worthy of further research, but our sense is that the net impact on the North Carolina (or, more broadly, the Southeastern) economy is minimal. Second, our analysis does not take into account benefits accruing to non-ORV users who derive more satisfaction from CAHA trips with fewer ORVs. Our informal discussions with several CAHA visitors suggest these benefits are real, and research by Mansfield et al. (2008) and Magee (2008) suggest they may be substantial. Finally, we do not include in our benefit calculations the WTP for endangered species protection in CAHA of residents outside of North Carolina. Careful accounting of these benefits would surely make the benefit-cost ratio even more favorable to the policy.

VII. Conclusion

Management of public land for multiple uses is a continual challenge for federal agencies. When management decisions involve competing non-market amenities, the economic benefits and costs may be difficult to assess. Here, we examine the impact of restricting access to coastal recreation to support endangered species protection. Our results suggest that the welfare losses to shoreline recreational anglers associated with access restrictions on CAHA are lower than the likely benefits for the public. The welfare cost estimates range from $403,000 to $2.07 million annually depending on the management strategy and level of closures in adaptive planning areas. The results also suggest that switching from the current management scheme to an alternative with greater focus on preservation of threatened species would have a marginal impact on welfare. Demand responses to the policies are relatively small and suggest that adaptation in recreation choice is likely to occur and offset some of the losses. It is important to note that the welfare losses estimated here only provide guidance on a portion of the economic impacts of the ORV restriction on CAHA. Yet, when we account for congestion, non-local recreation, enforcement, and other ORV-based recreation impacts, estimated benefits still outweigh costs.

In addition to the policy-relevant welfare results, our research also provides a methodological framework for estimating welfare effects of time-varying closures using the MRIP data – a large, publically available dataset – that incorporates two independent surveys to directly estimate participation, avoids potential bias associated with site aggregation, and addresses critiques of such a model by Train (2016). The methods here could be applied to value welfare losses related to any type of closure, including oil spills (see Alvarez et al. 2014), fish advisories, and beach closures due to dangerous levels of pollution.

Our results also come at a time when an additional federal statute, the 2015 NDAA, has forced the NPS to again revisit ORV management on CAHA. The policies have been highly contested over the past decade and this new legislation continues that debate. When ultimately resolved, the final outcome in CAHA is likely to have far-reaching impacts on shoreline management strategies in other NPS-administrated areas, such as Cape Lookout (NC) and Padre Island (TX) National Seashores. Overall, the new proposed CAHA revisions are likely to reduce closure times and increase access (i.e., reduce costs) with minimal impacts on biodiversity, further strengthening our findings that ORV restrictions are a beneficial policy measure.

References

Alvarez, S. S.L. Larkin, J.C. Whitehead, and T. Haab. 2014. "A revealed preference approach to valuing non-market recreational fishing losses from the Deepwater Horizon oil spill." *Journal of Environmental Management* 145: 199-209.

Ben-Akiva, M. and S.R. Lerman. 1985. *Discrete Choice Analysis*. Cambridge, MA: The MIT Press.

Berry, S.T. 1994. "Estimating discrete-choice models of product differentiation." *RAND Journal of Economics* 25:242-262.

Breidt, F. J., H.-L. Lai, J. D. Opsomer, and D. A. Van Voorhees. 2012. *A Report of the MRIP Sampling and Estimation Project: Improved Estimation Methods for the Access Point Angler Intercept Survey Component of the Marine Recreational Fishery Statistics Survey*. Silver Spring, MD: NOAA National Marine Fisheries Service, January.

Bujosa, A., A. Reira, R.L. Hicks, and K.E. McConnell. 2015. "Densities rather than shares: improving the measurement of congestion in recreation demand models." *Environmental and Resource Economics* 61: 127-140.

Cesario, F.J. 1976. "The Value of Time in Recreation Benefit Studies." *Land Economics* 52:32-41.

Dalrymple, C.J., M.N. Peterson, D.T. Cobb, E.O. Sills, H.D. Bondell, and D.J. Dalrymple. 2012. "Estimating Public Willingness to Fund Nongame Conservation Through State Tax Initiatives." *Wildlife Society Bulletin* 36:483-491.

Deisenroth, D.D, J.B. Loomis, and C.A. Bond. 2009. "Non-market valuation of off-highway vehicle recreation in Larimer County, Colorado: Implications of Trail Closures." *Journal of Environmental Management*, 90: 3490-97.

Englin, J., T. Holmes and R. Niell. 2006. "Alternative models of recreational off-highway vehicle site demand." *Environmental and Resource Economics*, 35: 327-338.

Haab, T.C. and K.E. McConnell. 2002. *Valuing Environmental and Natural Resources.* Northampton, MA: Edward Elger Publishing.

Hausman, J.A., G.K. Leonard and D. McFadden. 1995. "A Utility-Consistent, Combined Discrete Choice and Count Data Model Assessing Recreational Use Losses Due to Natural Resource Damage." *Journal of Public Economics* 56:1-30.

Herriges, J.A. and C.L. Kling. 1997. "The Performance of Nested Logit Models when Welfare Estimation is the Goal." *American Journal of Agricultural Economics* 79:792-802.

Hindsley, P., C.E. Landry, and B. Gentner. 2011. "Addressing onsite sampling in recreation site choice models." *Journal of Environmental Economics and Management* 62:95-110.

Jakus, P.M., J.E. Keith, L. Liu, and S. Blahna. 2010. "The Welfare Effects of Restricting Off-Highway Vehicle Access to Public Lands. *Agricultural and Resource Economics Review* 39: 89-100.

Johnston, R.J., and K. Moeltner. 2014. "Meta-modeling and Benefit Transfer: The Empirical Relevance of Source-consistency in Welfare Measures." *Environmental and Resource Economics* 59:337-361.

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

Landry, C.E., A.R. Lewis, H. Liu, and H. Vogelsong. 2016. "Addressing Onsite Sampling in Analysis of Recreation Demand: Economic Value and Impact of Visitation to Cape Hatteras National Seashore." *Marine Resource Economics* 31: 301 – 322.

Lovell, S.J. and D.W. Carter. 2014. "The use of sampling weights in regression models of recreation fishing-site choices." *Fishery Bulletin* 112:243-252.

Magee, L.E. 2008. "Closing Motor Vehicle Beach Access in the Mid-Atlantic: Implications for Social Welfare," MS thesis, University of Delaware.

Mansfield C., D.J. Phanuef, F.R. Johnson, J-C. Yang, and R. Beach. 2008. Preferences for Pubic Lands Management under Competing Uses: The Case of Yellowstone National Park. *Land Economics* 84:283-305.

Mansfield, C., R. Loomis, Evans, B. and B. Munoz. 2010. *Visitor Intercept Survey: Off-road Vehicle Management, Cape Hatteras National Seashore*. Denver: National Park Service, Final Report, October.

Mansfield, C., R. Loomis, and F. Braun. 2011. *Benefit-cost Analysis of Proposed ORV Use Regulation in Cape Hatteras National Seashore*. Denver: National Park Service, Final Report, January.

Moeltner, K., and R.S. Rosenberger. 2014. "Cross-context Benefit Transfer: a Bayesian Search for Information Pools." *American Journal of Agricultural Economics* 96:469-488.

Morey, E.R., R.D. Rowe, and M. Watson. 1993. "A Repeated Nested-logit Model of Atlantic Salmon Fishing." *American Journal of Agricultural Economics* 75:578-592.

Morey, E.R. 1999. "Two RUMs Uncloaked: Nested-logit Models of Site Choice and Nested-logit Models of Participation and Site Choice." In: Herriges, J.A., Kling, C.L., eds. *Valuing recreation and the environment: revealed preference methods in theory and practice*. Cheltenham, UK: Edward Elgar Publishing, Chapter 4.

National Park Service. 2010. *Cape Hatteras National Seashore Off-road Vehicle Management Plan: Final Environmental Impact Assessment*. Washington DC: November.

Parsons G.R. and M.S. Needelman. 1992. "Site aggregation in a random utility model of recreation." *Land Economics*, 68:418-433.

Stanley, D. L. 2005. "Local Perception of Public Goods: Recent Assessments of Willingness-to-Pay for Endangered Species." *Contemporary Economic Policy* 23:165–179.

Train, K. 2016. "Comment on "A revealed preference approach to valuing non-market recreational fishing losses from the Deepwater Horizon oil spill and its "Corrigendum" by Alvarez et al**.**" *Journal of Environmental Management* 167: 259-261.

Timmins, C. and J. Murdoch. 2007. "A revealed preference approach to the measurement of congestion in travel cost models." *Journal of Environmental Economics and Management* 53: 230-249.

USFW (U.S. Fish and Wildlife Service). 2011. Abundance and Productivity Estimates – 2011 Update, Atlantic Coast Piping Plover Population. Hadley, MA.

Whitehead, J.C. 1993. "Economic Values for Coastal and Marine Wildlife: Specification, Validity, and Valuation Issues." *Marine Resource Economics* 8:119-132.

Whitehead, J.C. and T.C. Haab. 1999. "Southeast marine recreational fishery statistical survey: distance and catch based choice sets. *Marine Resource Economics*, 14: 283-298.

Whitehead, J.C., B. Poulter, C. F. Dumas, and O. Bin. 2009. "Measuring the economic effects of sea level rise on shore fishing." *Mitigation and Adaptation Strategies for Global Change* 14: 777-792.

Williams, T. 2012. "The Battle Over a North Carolina Beach Continues." *Audubon*, September-October.

Table 1: Definition of Variables

	Description	Mean	Min	Max
Variables Entering Site Choice Model				
Site-Specific Travel Cost (TC)	From zip code of origin to each intercept site			
Alternative Specific Constants (ASCs)	ASCs for each site choice captures all site characteristics that are the same across individuals, both observed and unobserved			
Variables Entering Participation Model		Mean	Min	Max
Demographics (Phone Exchange level)				
Income	Average annual real household income	\$70,777	\$20,884	\$345,497
Population Density	People per square mile	5,650	0.11	130,031
White	% of population that is white	0.70	0.04	1
Male	% of population that is male	0.49	0.32	0.84
Education	% of population completing bachelor's degree or higher	0.30	0	0.91
CAHA Proximity Indicator	Identifies individuals residing > 300 mile one-way driving distance from CAHA			
No Trip ASC	ASC for no trip alternative			
Inclusive Value	Expected utility for the alternative choice			

Note: Participation model also includes area code and year fixed effects. Demographic variables are at the phone exchange level and represent a population weighted average of the zip code demographic data contained in each exchange. *Source:* Demographic data obtained from U.S. Census American Community Survey.

Table 2: Model Estimates

Panel A. Site Choice Model	*Parameter*	*T-stat*
Travel Cost		
2005	-0.073***	-10.74
2006	-0.072***	-9.20
2007	-0.075***	-12.15

Panel B. Participation Model	**Model 1**		**Model 2**		**Model 3**		**Model 4 (preferred)**	
	Parameter	*Std. Err.*	*Parameter*	*Std. Err.*	*Parameter*	*Std. Err.*	*Parameter*	*Std. Err.*
Constant	-6.329***	0.581	-6.338***	0.580	-5.631***	0.561	-9.719***	0.855
Dissimilarity Coefficient	0.060***	0.005	-	-	0.024***	0.009	0.036*	0.022
2005	-	-	0.056***	0.007	-	-	-	-
2006	-	-	0.067***	0.006	-	-	-	-
2007	-	-	0.059***	0.006	-	-	-	-
Demographics								
Average HH Income	-5.8e-06**	2.5e-06	-5.9e-06**	2.4e-06	7.0e-06***	2.3e-06	8.7e-06**	3.4e-06
Percent White	2.7e-06	2.9e-06	2.7e-06	2.9e-06	3.4e-06	3.0e-06	2.3e-07	3.1e-06
Percent Bachelor's Degree	1.044**	0.422	1.054**	0.421	0.828**	0.411	-0.196	0.587
Percent Male	3.212***	1.078	3.227***	1.077	3.062***	1.061	1.725	1.129
Population Density	-0.0001***	0.000	-0.0001***	0.000	-0.0001***	0.000	-0.00002	0.000
Proximity Indicator (300 mi.)	-	-	-	-	-0.619***	0.120	-0.706***	0.135
Year Fixed Effects	N		N		N		Y	
Area Code Fixed Effects	N		N		N		Y	
Observations	19,860		19,860		19,860		19,860	
Model Fit (Pseudo R-squared)	0.0181		0.0183		0.0211		0.0403	

Notes: All models are estimated conservatively with robust standard errors clustered by phone exchange.
*** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

Table 3: Off-Road Vehicle Restriction Policy Scenarios: *Alternative F*

Fishing Site	Island	*Wave 1*	*Wave 2*	*Wave 3*	*Wave 4*	*Wave 5*	*Wave 6*
Oregon Inlet (North)	Bodie	O	A	A	A	O	O
Rodanthe Fishing Pier	Hatteras	XP	XP	O	O	XP	XP
Beach Access Ramp 20	Hatteras	XP	XP	O	O	XP	XP
Beach Access Ramp 23	Hatteras	O	O	O	O	O	O
Beach Access Ramp 27	Hatteras	XP	XP	XP	XP	XP	XP
Beach Access Ramp 30	Hatteras	O	O	O	O	O	O
Beach Access Ramp 34	Hatteras	O	A	A	A	O	O
Avon Fishing Pier	Hatteras	XP	XP	O	O	XP	XP
Beach Access Ramp 38	Hatteras	O	O	O	O	O	O
Buxton Beach	Hatteras	XP	XP	O	O	XP	XP
Cape Point	Hatteras	O	A	A	A	O	O
Beach Access Ramp 49	Hatteras	O	O	O	O	O	O
Frisco Pier	Hatteras	XP	XP	O	O	XP	XP
Hatteras Inlet	Hatteras	X	A	A	A	A	X
Hatteras Inlet Beach	Ocracoke	X	A	A	A	A	X
Ocracoke Inlet & Beach	Ocracoke	O	A	A	A	A	O

Notes: O = Open, no impact. X = Closed: ORV restrictions and need ORV for access. XP = ORV restrictions but pedestrian access. A = Adaptive management with closures – could be O or X.

Table 4: Definition of Upper and Lower Bounds for Policy Simulations

Policy Scenarios	*Lower Bound*	*Upper Bound*
O	Site Open to ORVs	Site Open to ORVs
X	Site Closed to ORV & Pedestrian Access	Site Closed to ORV & Pedestrian Access
XP	Site Open to Pedestrian Access Only; Add 2 Hours of Travel Time [1]	Site Closed to ORV & Inaccessible to Pedestrians
A	Site Open to ORVs	Site Closed to ORV & Pedestrian Access

Notes: [1] – This represents additional cost of accessing a fishing site on foot instead of with an ORV.

Table 5: Welfare Costs Associated with Policy Scenarios

(in thousands of 2010$)	*Upper Bound*			*Lower Bound*		
Alternative F	Estimate	95% CI		Estimate	95% CI	
Aggregate	-$2,068	-$1,349	-$2,982	-$403	-$239	-$601
Year-Specific						
2005	-$1,445	-$1,291	-$1,707	-$254	-$227	-$299
2006	-$2,639	-$2,383	-$3,207	-$531	-$485	-$639
2007	-$2,121	-$1,877	-$2,506	-$423	-$378	-$496
Alternative D						
Aggregate	-$2,746	-$1,774	-$3,964	-$697	-$423	-$1,018
Year-Specific						
2005	-$1,899	-$1,696	-$2,242	-$451	-$403	-$529
2006	-$3,509	-$3,171	-$4,260	-$899	-$821	-$1,082
2007	-$2,830	-$2,507	-$3,340	-$741	-$662	-$869
Close all CAHA sites						
Aggregate	-$3,603	-$2,286	-$5,240			
Year-Specific						
2005	-$2,448	-$2,186	-$2,893			
2006	-$4,638	-$4,186	-$5,635			
2007	-$3,724	-$3,294	-$4,399			

Notes: All numbers are in thousands of 2010 US dollars. Models are calibrated to impose a dissimilarity coefficient (0.46) and imputed value of a trip ($30) supported by recent meta-analyses [Johnston and Moeltner (2014) and Moeltner and Rosenberger (2014)]. Simulation estimates are for WTP (in 2010 dollars) for residents of coastal counties covered by the MRIP survey. The *upper bound* of welfare is estimated given the most restrictive possibilities on ORV rules. The *lower bound* of welfare is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).

Table 6: Demand Responses for Policy Scenarios

(in thousands of trips)	*Upper Bound*			*Lower Bound*		
Alternative F	Estimate	95% CI		Estimate	95% CI	
Affected Trips	143	92	197	143	92	197
Lost	68	44	96	13	7.8	19
Substitute	76	48	102	14	8.2	19
Diminished	-	-	-	116	76	159
Alternative D						
Affected Trips	187	120	256	187	120	256
Lost	90	58	127	23	14	33
Substitute	97	62	131	23	14	30
Diminished	-	-	-	141	92	194
Close all CAHA sites						
Affected Trips	239	152	329			
Lost	118	75	168			
Substitute	121	77	163			
Diminished	-	-	-			

Notes: All numbers are in thousands of trips. Models are calibrated to impose a dissimilarity coefficient (0.46) and imputed value of a trip ($30) supported by recent meta-analyses [Johnston and Moeltner (2014) and Moeltner and Rosenberger (2014)]. The *upper bound* is estimated given the most restrictive possibilities on ORV rules. The *lower bound* is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).

Table 7: Aggregate *Alternative F* Welfare Predictions with Different Model Specifications

Model Specification	*Upper Bound*				*Lower Bound*			
(in thousands of 2010$)	Estimate	95% CI		t-Stat	Estimate	95% CI		t-Stat
Calibrated Site/Wave Choice	-$2,068	-$1,349	-$2,982	-3.93	-$403	-$239	-$601	-3.34
Calibrated Site Choice Only	-$2,071	-$1,479	-$2,790	-4.99	-$440	-$324	-$583	-5.58
Uncalibrated Site/Wave Choice	-$2,163	-$1,405	-$3,129	-3.87	-$406	-$240	-$605	-3.33
Uncalibrated Site choice Only	-$2,177	-$1,547	-$2,943	-4.89	-$446	-$328	-$590	-5.57

Notes: All numbers are in thousands of 2010 US dollars. Models are calibrated to impose a dissimilarity coefficient (0.46) and imputed value of a trip ($30) supported by recent meta-analyses [Johnston and Moeltner (2014) and Moeltner and Rosenberger (2014)]. Simulation estimates are for WTP (in 2010 dollars) for residents of coastal counties covered by the MRIP survey. The *upper bound* of welfare is estimated given the most restrictive possibilities on ORV rules. The *lower bound* of welfare is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).
Tables with full results similar to Table 5 for 'Calibrated Site Choice Only' and 'Uncalibrated Site/Wave Choice' are provided in Appendix B.

Table 8: Comparison of Local and Non-Local Anglers at Three Spatial Scales

Trip Location	**Three-State Region** *Mean*	**North Carolina** *Mean*	**Dare County** *Mean*
Coastal County Origin			
Hours Fished	3.79	3.69	3.99
Catch (Binary)	0.20	0.17	0.15
Fish by Individual	6.87	7.02	7.17
Mode of Fishing			
Beach	0.41	0.54	0.51
Pier	0.53	0.42	0.38
Observations	10,298,584	6,883,154	3,762,994
Non-Coastal Origin			
Hours Fished	3.73	3.69	3.82
Catch (Binary)	0.14	0.14	0.11
Fish by Individual	7.17	7.22	7.42
Mode of Fishing			
Beach	0.35	0.49	0.57
Pier	0.59	0.47	0.44
Observations	7,931,929	5,231,832	3,353,770

Source: Authors calculations from MRIP survey data.

Table 9: Upper Bound of Potential Additional Welfare Costs under *Alternative F*

Model Estimates	Upper Bound: **$2.07 Million**	Lower Bound: **$403,000**
Limitations	*Additional Costs*	
Congestion	+ $1.04 Million	+ $202,000
Non-local Recreation	+ $2.65 Million	+ $515,000
Other ORV Recreation	+ $5.76 Million	+ $1.12 Million
Enforcement Costs	+ $1.10 Million	+ $1.10 Million
Maximum Potential Costs	**$12.62 Million**	**$3.34 Million**

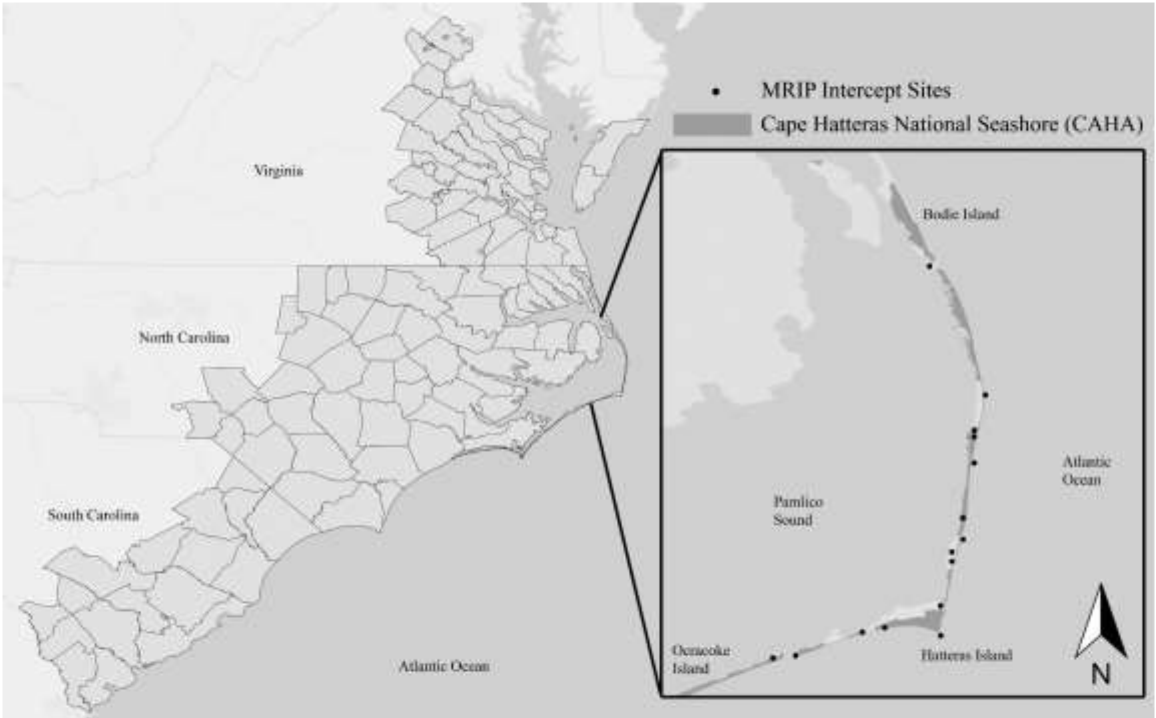

Figure 1: Study Area and MRIP Intercept Site Locations

Note: Coastal counties included in participation model are highlighted in light grey. Cape Hatteras National Seashore is highlighted in dark grey in the inset and the black dots indicate the location of 16 MRIP intercept sites potentially impacted by the NPS management policies.

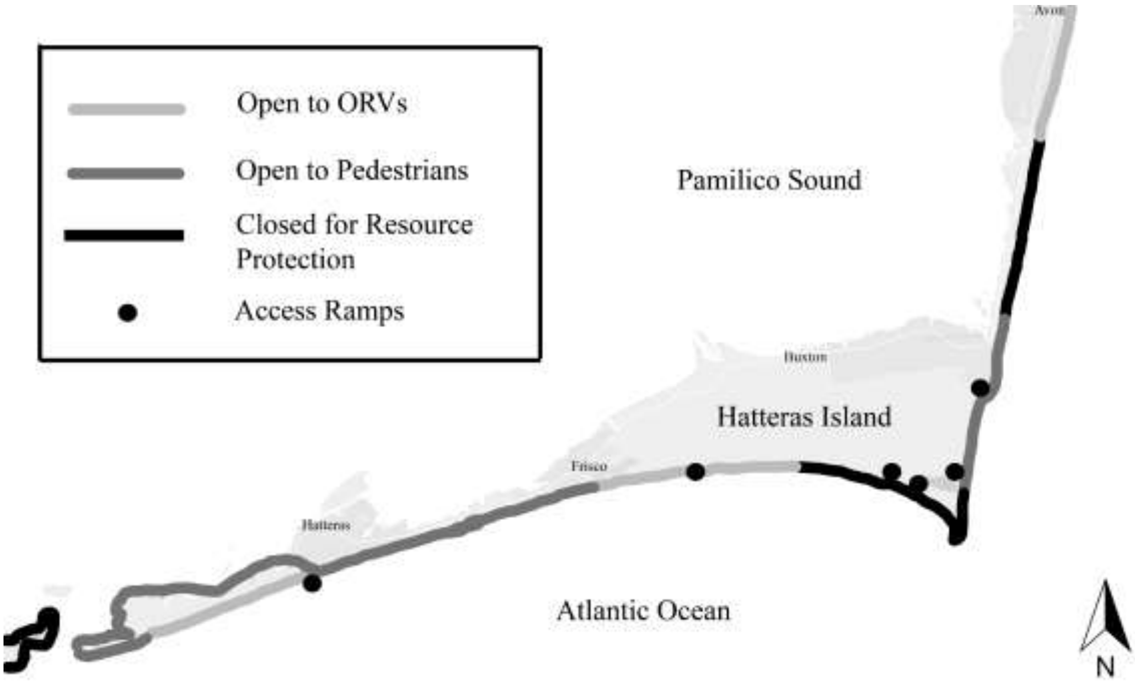

Figure 2: ORV Restrictions on Hatteras Island: June 1, 2015

Source: National Park Service website

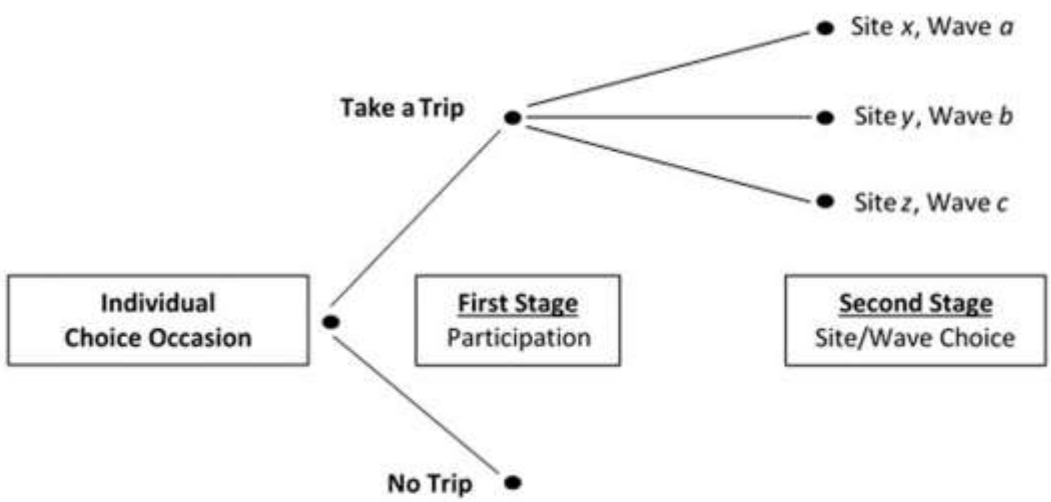

Figure 3: Example Decision Tree for Two-Level Nested Logit Model

Note: This is a general example of a decision tree and does not represent a full enumeration of the site-wave choices anglers face in our models.

APPENDIX

Appendix A: Site Registry Data

This appendix describes the steps used to transform the trip frequency information contained in the MRIP site registries into aggregate trip estimates for each of the 344 shoreline fishing sites in the three-state region (NC, SC, and VA). Every two months, NOAA updates its master list of public access shoreline fishing sites, or site registry. Each site has weekday and weekend trip frequency or "fishing pressure" estimates associated with it, which are also updated bimonthly. These estimates represent NOAA's best estimate of the number of trips occurring at a site in a normal 8-hour period, and this information informs whether and how intensely to sample at each site in each wave. Bimonthly updates are based on feedback from NOAA field staff as well as auxiliary sources (e.g., published newspaper reports about pier closures).

To construct estimates of aggregate trips for each MRIP site/wave combination, we employ the following steps. First, weekday and weekend trip estimates are constructed for each site and wave from the contemporaneous site registry. We assume that the average fishing day is 16 hours at manmade sites (e.g., piers) which are generally lighted and 12 hours for all other sites. These daily estimates are then aggregated to the bimonthly period. Finally, a regression-based adjustment is made to these estimates to account for the fact that not all trips at a site originate from coastal counties. Data from the MRIP intercept survey is used for this task. Specifically, intercepted respondents report their home zip code, which allows us to determine if they live in a coastal or non-coastal county. For every sampled site, the share of trips originating from coastal counties can be constructed, and because sampling is by design simple random sampling at the site level, this constructed share is an unbiased estimate of the population share at that site. A weighted linear regression is then used to predict the share of coastal trips as a function of observable site characteristics. The weights employed in the regression analysis are inversely proportional to the intensity of sampling at the sites. These predicted shares are then combined with the bimonthly trip estimates to generate total trip predictions from coastal counties for all 344 MRIP sites.

Appendix B: Additional Tables

Table B.1: Off-Road Vehicle Restriction Policy Scenarios: *Alternative D*

Fishing Site	Island	*Wave 1*	*Wave 2*	*Wave 3*	*Wave 4*	*Wave 5*	*Wave 6*
Oregon Inlet (North)	Bodie	X	A	A	A	A	X
Rodanthe Fishing Pier	Hatteras	XP	XP	O	O	XP	XP
Beach Access Ramp 20	Hatteras	XP	XP	O	O	XP	XP
Beach Access Ramp 23	Hatteras	O	O	O	O	O	O
Beach Access Ramp 27	Hatteras	XP	XP	XP	XP	XP	XP
Beach Access Ramp 30	Hatteras	A	A	A	A	A	A
Beach Access Ramp 34	Hatteras	XP	XP	XP	XP	XP	XP
Avon Fishing Pier	Hatteras	XP	XP	O	O	XP	XP
Beach Access Ramp 38	Hatteras	O	O	O	O	O	O
Buxton Beach	Hatteras	XP	XP	O	O	XP	XP
Cape Point	Hatteras	X	A	A	A	A	X
Beach Access Ramp 49	Hatteras	O	O	O	O	O	O
Frisco Pier	Hatteras	XP	XP	O	O	XP	XP
Hatteras Inlet	Hatteras	X	A	A	A	A	X
Hatteras Inlet Beach	Ocracoke	X	A	A	A	A	X
Ocracoke Inlet & Beach	Ocracoke	X	A	A	A	A	X

Notes: O = Open, no impact. X = Closed: ORV restrictions and need ORV for access. XP = ORV restrictions but pedestrian access. A = Adaptive management with closures – could be O or X.

Table B.2: Welfare Costs Associated with Policy Scenarios (Uncalibrated)

(in thousands of 2010$)	*Upper Bound*			*Lower Bound*		
Alternative F	Estimate	95% CI		Estimate	95% CI	
Aggregate	-$2,163	-$1,405	-$3,128	-$406	-$240	-$605
Year-Specific						
2005	-$1,501	-$1,345	-$1,777	-$255	-$229	-$302
2006	-$2,768	-$2,503	-$3,340	-$535	-$489	-$639
2007	-$2,219	-$1,962	-$2,623	-$426	-$381	-$499
Alternative D						
Aggregate	-$2,920	-$1,870	-$4,230	-$706	-$428	-$1,032
Year-Specific						
2005	-$1,995	-$1,787	-$2,363	-$454	-$408	-$537
2006	-$3,742	-$3,387	-$4,512	-$911	-$833	-$1,088
2007	-$3,007	-$2,660	-$3,556	-$750	-$671	-$879
Close all CAHA sites						
Aggregate	-$3,927	-$2,462	-$5,753			
Year-Specific						
2005	-$2,629	-$2,353	-$3,122			
2006	-$5,089	-$4,605	-$6,142			
2007	-$4,063	-$3,585	-$4,822			

Notes: All numbers are in thousands of 2010 US dollars. Models allow uncalibrated dissimilarity coefficient (0.04) and imputed value of a trip ($342). Simulation estimates are for WTP (in 2010 dollars) for residents of coastal counties covered by the MRIP survey. The *upper bound* of welfare is estimated given the most restrictive possibilities on ORV rules. The *lower bound* of welfare is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).

Table B.3: Demand Responses for Policy Scenarios (Uncalibrated)

(in thousands of trips)	*Upper Bound*			*Lower Bound*		
Alternative F	Estimate	95% CI		Estimate	95% CI	
Affected Trips	143	92	197	143	92	197
Lost	5.7	-0.4	15	1.1	-0.0	2.9
Substitute	138	88	192	24	14	34
Diminished	-	-	-	118	77	161
Alternative D						
Affected Trips	187	120	256	187	120	256
Lost	7.6	-0.5	20	1.8	-0.1	5.0
Substitute	179	114	250	40	25	55
Diminished	-	-	-	145	94	199
Close all CAHA sites						
Affected Trips	239	152	329			
Lost	10	-0.7	27			
Substitute	229	144	321			
Diminished	-	-	-			

Notes: All numbers are in thousands of trips. Models allow uncalibrated dissimilarity coefficient (0.04) and imputed value of a trip ($342). The *upper bound* is estimated given the most restrictive possibilities on ORV rules. The *lower bound* is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).

Table B.4: Welfare Costs Associated with Policy Scenarios (Site Choice Only)

(in thousands of 2010$)	*Upper Bound*			*Lower Bound*		
Alternative F	Estimate	95% CI		Estimate	95% CI	
Aggregate	-$2,171	-$1,479	-$2,790	-$440	-$324	-$583
Year-Specific						
2005	-$1,548	-$1,431	-$1,777	-$347	-$319	-$404
2006	-$2,488	-$2,323	-$2,911	-$524	-$493	-$600
2007	-$2,177	-$2,029	-$2,490	-$450	-$404	-$528
Alternative D						
Aggregate	-$2,820	-$2,104	-$3,715	-$730	-$559	-$969
Year-Specific						
2005	-$2,216	-$2,040	-$2,563	-$597	-$541	-$702
2006	-$3,334	-$3,142	-$3,867	-$874	-$815	-$1,016
2007	-$2,912	-$2,659	-$3,393	-$721	-$632	-$862
Close all CAHA sites						
Aggregate	-$3,622	-$2,610	-$4,835			
Year-Specific						
2005	-$2,747	-$2,525	-$3,174			
2006	-$4,338	-$4,087	-$5,043			
2007	-$3,782	-$3,461	-$4,387			

Notes: This alternative model specification where each individual's choice set consists of sites as opposed to site/wave pair. All numbers are in thousands of 2010 US dollars. Model is calibrated to impose a dissimilarity coefficient (0.46) and imputed value of a trip ($30) supported by recent meta-analyses [Johnston and Moeltner (2014) and Moeltner and Rosenberger (2014)]. Simulation estimates are for WTP (in 2010 dollars) for residents of coastal counties covered by the MRIP survey. The *upper bound* of welfare is estimated given the most restrictive possibilities on ORV rules. The *lower bound* of welfare is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).

Table B.5: Demand Responses for Policy Scenarios (Site Choice Only)

(in thousands of trips)	*Upper Bound*			*Lower Bound*		
Alternative F	Estimate	95% CI		Estimate	95% CI	
Affected Trips	146	98	189	146	98	189
Lost	69	47	91	15	10	19
Substitute	138	88	192	24	14	34
Diminished	-	-	-	116	77	151
Alternative D						
Affected Trips	195	138	248	195	138	248
Lost	93	66	121	24	18	32
Substitute	101	72	129	24	18	31
Diminished	-	-	-	147	102	188
Close all CAHA sites						
Affected Trips	244	167	314			
Lost	130	82	157			
Substitute	124	86	158			
Diminished	-	-	-			

Notes: This alternative model specification where each individual's choice set consists of sites as opposed to site/wave pair. All numbers are in thousands of trips. Model is calibrated to impose a dissimilarity coefficient (0.46) and imputed value of a trip ($30) supported by recent meta-analyses [Johnston and Moeltner (2014) and Moeltner and Rosenberger (2014)]. The *upper bound* is estimated given the most restrictive possibilities on ORV rules. The *lower bound* is estimated given the most relaxed possibilities on ORV rules. **Alternative F** was implemented by NPS and two stricter scenarios, **Alternative D** and **Close all CAHA sites,** are shown for comparison purposes. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986).