Recreation Costs of Endangered Species Protection: Evidence from Cape Hatteras National Seashore

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Abstract: Management of public lands often involves competing uses and difficult tradeoffs. In this article, we examine the impact of an economically important, policy-relevant public land management regulation designed to protect coastal biodiversity. We focus our attention on the land use conflict at Cape Hatteras National Seashore between off-road vehicle (ORV) access and nesting habitat protection for a number of endangered species. We combine site choice and participation data to estimate a repeated discrete choice model of recreational angler behavior in response to time-varying access restrictions. Our results suggest the economic costs of this policy 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 an upper bound of recreation costs is likely less than conservative estimates of the benefits associated with endangered species protection.

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INTRODUCTION

Current public land holdings by the United States government span 640 million acres, or more than one quarter of all US land. The Bureau of Land Management (BLM) and the US 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 (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, at 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 at 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, Loomis, and Bond (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, Holmes, and Niell (2006) and Jakus et al. (2010) also investigate the costs of ORV closures on public lands, yet only the latter attempts to monetize the welfare costs from trail closures. In their study of Utah federal lands, Jakus et al. (2010) find costs on the order of $1.2 million annually.

Herein, 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 pre-ORV policy 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.[3] Our models are 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 NPS's 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%.[4] This finding suggests that most anglers will either continue to take fishing trips to CAHA 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).

The welfare losses estimated in our models may not represent all potential costs associated with the NPS policy. To provide policy-relevant context building on our results, we then construct an upper bound of potential costs, including: (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. We combine our results with previous findings in the literature, CAHA visitor survey results, and anecdotal evidence from the NPS to conservatively inflate the costs of the ORV restrictions to obtain a minimum benefit-cost ratio for the final welfare comparison. In this exercise, the losses from CAHA ORV restrictions could 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 contingent valuation estimates in Whitehead (1993) ($13 million annually) and Dalrymple et al. (2012) ($48 million annually). Given this information, it is likely that current ORV restrictions in CAHA pass a benefit-cost test, although refinements to current policies could generate higher net benefits.

In addition 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 effort. 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, Landry, and Gentner 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, marinas), 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 closures were contingent on park managers discovering nesting sites of endangered or threatened species, NPS announced well in advance that adaptive management plans were in place and could lead to closures.[5] Therefore, when anglers are planning when and where to recreate each year, they can anticipate closures that will or may be in place at different times and places and respond accordingly. This predictability and adaptability mitigates the "time travel" 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.

CONFLICT AT CAPE HATTERS NATIONAL SEASHORE

CAHA was authorized by Congress in 1937 and established in 1953 as the first national seashore in the US. The barrier island park stretches over 67 miles, contains three islands (Hatteras, Ocracoke, and Bodie), and covers 24,470 acres of North Carolina's coastline (see figure 1).[6] 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 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 experienced declines in populations of plant and animal species, ranging from nesting shorebirds (e.g., roseate tern) to small beach plants (e.g., seabeach amaranth). The charismatic species of concern in CAHA affected by ORV use that generate the most attention are the piping plover (*Charadrius melodus*) and sea turtles. This plover is a protected species under the federal Endangered Species Act (ESA) and is designated as threatened.[7] Piping plover populations are vulnerable to habitat loss resulting 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*).[8] 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 both species, North Carolina represents the northern limits of their nesting grounds.[9] 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 ESA, 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 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.[10]

The final management rule went into effect on February 15, 2012, after four years of policy uncertainty. The NPS chose between a number of alternative plans after a period of public comment (National Park Service 2010; Mansfield, Loomis, and Braun 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. 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*) 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* designated 27.2 miles for year-round ORV routes, 0 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. 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; and (3) limiting vehicle-free areas. After five public hearings in August 2015 and further internal deliberations, NPS published proposed revisions to the CAHA management plan in August 2016 and, after a public comment period, final revisions in December 2016 in the *Federal Register*. Although we consider only the existing ORV rules in our policy simulations, the revisions were relatively modest in scale and will, if anything, reduce regulatory costs and strengthen our overall finding that the net-benefits of the ORV restrictions are positive.[11]

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, Rowe, and Watson 1993) to model whether, where, and *when* anglers engage in shoreline recreational fishing. An individual first chooses whether to participate in coastal fishing in a given year and then conditionally at which site, as well as in 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 for a different time of year when that 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 the "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, individuals choose the alternative that maximizes their 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:

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

where m_{it} is income; c_{ijw} is travel cost[12]; $\mathbf{X}_{jwt}$ is a vector of site characteristics; and ε_{ijwt} captures idiosyncratic, random factors. Conditional on taking a trip, a rational angler selects the site and wave 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, 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; 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 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.[13]

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

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

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:

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

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 commonly experienced by all individuals who visit a site 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 $\left(\beta / \lambda\right)$. However, since the MRIP survey samples, at most, 40 % of fishing sites in each wave/year, this step does not generate consistent estimates for the ASCs because they are not identified for unsampled sites. Therefore, our second step uses auxiliary data on aggregate trip frequency at every site to calibrate all ASCs. We 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. This auxiliary data 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 eight-hour period at the time sampling commences.[14] These estimates reflect visitation rates for individuals living inside and outside coastal counties. Bimonthly updates are based on feedback from NOAA field staff, as well as auxiliary sources (e.g., published newspaper reports about pier closures).

To generate visitation estimates for 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. 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 originate from coastal counties. 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. 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.

The adjusted ASCs are then used to generate the inclusive value index:

$$IV_i = \ln\left(\sum_{w=1}^{W}\sum_{k=1}^{J} e^{\left(\frac{\widehat{\beta}}{\lambda}c_{ikw}+\tilde{\delta}_{kw}\right)}\right), \qquad (6)$$

where $\frac{\widehat{\beta}}{\lambda}$ 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, Leonard, and McFadden 1995).

The third and final step estimates a standard discrete choice logit model of participation as a function of the inclusive value, demographics, a proximity to CAHA indicator,[15] area code fixed effects, and year fixed effects. We assume the total number of choice occasions on which an individual may take a trip is proportional to the length of the reporting period (i.e., two-month wave), scaled by an adjustment factor that ensures all respondents have more choice occasions than trips. For example, if the most avid angler reports taking 122 trips in a 61-day wave, the adjustment factor would equal two to ensure that all individuals have at least enough choice occasions to rationalize their trips. Although not reported here, sensitivity analysis suggested that our welfare results changed by less than 1% when we adopted alternative approaches to specify the adjustment.

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; δ_0 is the ASC for the no-trip alternative; 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 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.

DATA

MRIP data from NOAA’s National Marine Fisheries Service (NMFS) were collected with the point-of-access Angler Intercept Survey and the Coastal Household Telephone Survey from 2005–2007. This timeframe is chosen to be representative of angler behavior in the region *before* new ORV restrictions were implemented. Our primary objective is to model pre-policy angler behavior and simulate how that behavior could change in response to the proposed NPS ORV closure scenarios. Since nighttime driving restrictions were put in place in 2008 and there was significant uncertainty about the policy prior to full implementation in February 2012 (i.e., which policy alternative and when would it start), the 2005–2007 data provides the best option to model pre-policy angler behavior.[16] We use our model results with this data to then simulate how angler behavior would change based on NPS resource closures from most restrictive to least restrictive under two different management scenarios (*Alternatives F* and *D*). In other words, we impose realistic closure simulations of the actual policy on pre-policy behavior to estimate the upper bound of potential welfare losses.

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. 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 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. 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. Across the 18 waves spanning 2005–2007, 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.

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 trip intercepts 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, Landry, and Gentner 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 US 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 US 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 ($\bar{n}$).[17] The opportunity cost of time (*oppc*) is determined using the common assumption of one-third 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:

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

Phone survey data are collected by county-stratified random-digit-dial (RDD) from coastal households from 2005–2007 on the frequency of fishing trips in the preceding two months. 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. We utilize data from coastal counties in the three-state region surrounding the Outer Banks of North Carolina with 4,928 six-digit phone exchanges as the spatial unit of analysis.[18] Survey records for individuals who actively participated in coastal recreational fishing are used to characterize participation in the model. However, the non-participating households are identified at the county level only–a less refined spatial scale than 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 is 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 dummies, 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 US Census Bureau's American Community Survey. Zip codes are linked to six-digit phone exchanges using a proprietary data set from *Melissa Data*.[19] A population-weighted average of zip code demographic data is assigned to each phone exchange.

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 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.[20] 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 to similar estimates in the literature at large. The implication here is a relatively large per-trip value of approximately $342, compared to results from two recent meta-analyses that included saltwater fishing trips.[21] Moeltner and Rosenberger (2014) report the average willingness to pay (WTP)/day for a saltwater fishing trip in the Northeast as $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 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, the welfare implications are robust to either approach, but angler behavior (i.e. lost trip or a substitute trip) is sensitive to the model specification.

Policy Simulations

We now investigate the welfare and behavioral 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, which 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. The simulated closures represent how the policies would be implemented by the NPS at each MRIP site under each alternative. 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.[22] 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 are prohibited. ‘XP’ indicates that ORVs are prohibited, but the site is still open to pedestrians. Lastly, ‘A’ indicates adaptive planning areas where a site could either be open (‘O’) or 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 a 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 is 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 the 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 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.[23] ASCs and inclusive values (equation 6) 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 the 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 500 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 and t-statistics 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. We define an affected trip as a trip that occurs to the affected sites (i.e., those impacted by the ORV restrictions) in the baseline scenario. There are three types of affected trips. A lost trip is a trip that occurs in the baseline scenario but does not happen under the ORV closure scenarios, measured as the change in region-wide trips. Alternatively, anglers can take a diminished trip to an affected site that may be closed to ORVs but still open to pedestrian traffic if they are willing to bear the additional costs of accessing the affected site by foot. Restated, a diminished trip is a trip to an affected site that remains open to pedestrians. Finally, individuals have the option to take a substitute trip to other locations unaffected by the management policy. For each year, the model is utilized to estimate the number of affected trips, lost trips, and diminished trips 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:

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

Approximately 143,000 (7.5% 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% 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%, or an additional loss of 50,000 trips. Of the total affected trips, 75 to 81% are diminished trips in the lower-bound simulations. By construction, there are zero diminished trips at the upper bound, as all affected sites are closed. At the lower bound, substitute trips represent approximately 10 to 12% of affected trips, with the aggregate numbers being relatively similar to the number 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% and 12% of the total for *Alternatives F* and *D*, respectively.

For robustness checks on our preferred model (calibrated dissimilarity coefficient and allowing intertemporal substitution), we estimate a number of model variations. These include a site/wave choice model with the uncalibrated dissimilarity coefficient and site choice only models (i.e., no intertemporal substitution), with and without the calibrated dissimilarity coefficient. The welfare results from these models under *Alternative F* are displayed in table 7. The losses are 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 to 5.3% (0.7 to 10.7%) of the preferred model estimates. These results lend support to our preferred model choice which captures both spatial and temporal substitution behavior and the calibration decision to account for measurement error in our survey data. It is important to note that our model choice does matter for predicted angler behavior, implying that our decomposition of affected trips into lost, diminished, and substitute trips from our preferred specification should be interpreted with caution. As expected, the uncalibrated model with a high implied trip value generates more substitute and diminished trips and fewer lost trips than the calibrated model. Full simulation results using each alternative model specification are provided in the online-only appendix tables A.2–A.7.

Potential for Additional Costs

This section identifies four factors that have potential to increase our main welfare results 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 factors and our subsequent assumptions to estimate an upper bound, or "worst-case" scenario, of the potential costs of CAHA ORV restrictions. This cost inflation exercise is used to compare the maximum of potential costs of the policy to a conservative lower bound estimate of benefits of protecting coastal biodiversity, the aim of the ORV restrictions.

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% or more. Bujosa et al. (2015) take a site density approach to improve congestion measurement and also find increases in costs of approximately 50% when site congestion levels are adjusted after closures. Therefore, we add a 50% increase in our cost estimates to account for the potential adverse effects of congestion on utility 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% of the user days to sites in the Outer Banks of North Carolina. This implies that 46% of the user days are comprised of 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 are relatively similar between these groups at varying spatial scales–three-state region, North Carolina, and Dare County, NC (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 for generating an upper-bound welfare estimate. 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% 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%. 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 NPS 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 four-mile 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 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.

DISCUSSION

To put the welfare loss estimates from our analysis and the back-of-the-envelope calculation described above into perspective, it is important to consider the benefits of 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 WTP of $10.98 for loggerhead sea turtle preservation and $14.74 for all non-game endangered and threatened species in coastal North Carolina (both in 1991 dollars).[24] Dalrymple et al. (2012) find household WTP for non-game endangered species ranging from $65 to $98.80 per year.

Aggregating these estimates to all 3.7 million North Carolina households and adjusting to 2010 dollars produces aggregate benefits from Whitehead of $65 to $87.3 million annually and $240 to $366 million annually from Dalrymple et al. According to NC Wildlife Resources Commission's seaturtle.org (September 30, 2015) website, approximately 20% 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, a long-run *minimum* benefit cost ratio of the ORV restrictions ranges from 1.03 to 3.8.

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 may very well be substantial. But from an aggregate 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.[25] 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. Furthermore, annual recreation visits at CAHA remain in excess of 2 million people after the ORV restrictions were implemented. 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 WTP for endangered species protection in CAHA for residents outside of North Carolina in our benefit calculations. Careful accounting of these benefits would surely make the benefit-cost ratio even more favorable to the policy.

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 likely 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[26]–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 (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, forced the NPS to revisit ORV management in CAHA. The recently completed regulatory review, which resulted in modest policy changes that, if anything, strengthen our findings about the overall net benefits of ORV restrictions, continued public debate about the inherent tradeoffs between recreational access and environmental preservation. Going forward, these discussions will continue as the NPS considers similar shoreline regulations on other lands it manages, such as Cape Lookout (NC) and Padre Island (TX) National Seashores.

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Table 1. Definition of Variables

	Description	Mean	Min.	Max.
Panel A: 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 identical across individuals, both observed and unobserved			
Panel B: Variables Entering Participation Model				
Demographics (phone exchange level)		Mean	Min.	Max.
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 US Census American Community Survey (https://www.census.gov/programs-surveys/acs/).

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	

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

Note: 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

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

Table 5. Welfare Costs Associated with Policy Scenarios

(Thousands of 2010$)	*Upper Bound*				*Lower Bound*			
Alternative F	Estimate	95% Confidence Interval		t-Stat	Estimate	95% Confidence Interval		t-Stat
Aggregate	-$2,068	-$1,347	-$2,937	-4.01	-$403	-$238	-$590	-3.39
Year-Specific								
2005	-$1,445	-$1,304	-$1,686	-15.0	-$254	-$229	-$296	-15.5
2006	-$2,639	-$2,398	-$3,032	-16.6	-$531	-$487	-$610	-17.4
2007	-$2,121	-$1,910	-$2,436	-15.8	-$423	-$382	-$486	-15.8
Alternative D								
Aggregate	-$2,746	-$1,769	-$3,902	-3.95	-$697	-$422	-$999	-3.58
Year-Specific								
2005	-$1,899	-$1,713	-$2,216	-15.1	-$451	-$405	-$525	-15.4
2006	-$3,509	-$3,192	-$4,032	-16.7	-$899	-$824	-$1,033	-17.4
2007	-$2,830	-$2,545	-$3,250	-15.7	-$741	-$668	-$852	-15.8
Close all CAHA Sites								
Aggregate	-$3,603	-$2,279	-$5,157	-3.82				
Year-Specific								
2005	-$2,448	-$2,207	-$2,857	-15.1				
2006	-$4,638	-$4,216	-$5,326	-16.7				
2007	-$3,724	-$3,344	-$4,277	-15.6				

Note: 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; Moeltner and Rosenberger 2014). Simulation estimates are for WTP 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) with 500 draws.

Table 6. Demand Responses for Policy Scenarios

(Thousands of Trips)	*Upper Bound*				*Lower Bound*			
Alternative F	Estimate	95% Confidence Interval		t-Stat	Estimate	95% Confidence Interval		t-Stat
Affected Trips	143	92	195	4.26	143	92	195	4.26
Lost	-68	-44	-94	-4.14	-13	-7.8	-19	-3.47
Substitute	76	48	102	4.37	14	8.1	19	3.70
Diminished	-	-	-	-	116	76	157	4.45
Alternative D								
Affected Trips	187	119	255	4.21	187	119	255	4.21
Lost	-90	-58	-125	-4.06	-23	-14	-32	-3.66
Substitute	97	62	130	4.36	23	14	31	4.01
Diminished	-	-	-	-	141	92	192	4.35
Close all CAHA Sites								
Affected Trips	239	151	327	4.12				
Lost	-118	-75	-165	-3.92				
Substitute	121	77	162	4.33				
Diminished	-	-	-	-				

Note: All numbers are in thousands of trips. Models are calibrated to impose a dissimilarity coefficient (0.46) and an imputed value of a trip ($30) supported by recent meta-analyses (Johnston and Moeltner 2014; 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) with 500 draws.

Table 7. Aggregate *Alternative F* Welfare and Behavioral Predictions with Different Model Specifications

Model Specification	*Upper Bound*				*Lower Bound*			
(Thousands of 2010$)	Estimate	95% CI		t-Stat	Estimate	95% CI		t-Stat
Calibrated λ Site/Wave Choice (preferred specification)	-$2,068	-$1,346	-$2,941	-4.01	-$403	-$238	-$591	-3.39
Calibrated λ Site Choice Only	-$2,071	-$1,442	-$2,781	-4.96	-$440	-$322	-$587	-5.55
Uncalibrated λ Site/Wave Choice	-$2,163	-$1,398	-$3,087	-3.95	-$406	-$239	-$594	-3.38
Uncalibrated λ Site Choice Only	-$2,177	-$1,505	-$2,941	-4.86	-$446	-$325	-$593	-5.53

Note: All numbers are in thousands of 2010 US dollars. Calibrated models are calibrated to impose a dissimilarity coefficient (0.46) and an imputed value of a trip ($30) supported by recent meta-analyses (Johnston and Moeltner 2014; Moeltner and Rosenberger 2014). Uncalibrated models allow an uncalibrated dissimilarity coefficient (0.04) and an imputed value of a trip ($342). Simulation estimates are for WTP 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. Confidence intervals are estimated using a parametric bootstrap (Krinsky and Robb 1986) with 500 draws. Tables with full results similar to table 5 for all other model specifications are provided in the online-only appendix.

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: Author 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 Titles (Figures included at 600 DPI in separate files)

Figure 1. Study Area and MRIP Intercept Site Locations

Note: Coastal counties included in the 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: NPS 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.

Endnotes

[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 overestimated, as the stated preference model used by Deisenroth, Loomis, and Bond (2009) does not allow for site substitution by recreators if particular trails are closed.

[3] The MRIP data from 2005-2007 are utilized primarily to avoid the potential confounds of changes to ORV policy on CAHA that began with nighttime driving restrictions in 2008 and the uncertainty about policy changes in the intervening years until full implementation of the policy in February 2012. This choice allows us to capture angler behavior before any policies were in place and estimate the welfare changes using NPS management alternatives at the site level.

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

[5] 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 when a change takes place.

[6] The US 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.

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

[8] 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.

[9] 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.

[10] This restriction appears to be helping, as the NPS has found an average nesting total in CAHA around 129 annually from 2008–2011.

[11] In particular the revisions will result in greater ORV access along three dimensions: 1) certain popular beaches will open earlier in the day; 2) beaches near the villages of Rodanthe, Waves, Salvo, Avon, Frisco and Hatteras will open two weeks earlier in the fall and close two weeks later in the spring; and 3) restrictions will be partially relaxed at some beaches that were previously designated as vehicle free. Overall, NPS anticipates that these restrictions will modestly lower regulatory costs and not significantly impact environmental benefits.

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

[13] 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.

[14] 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 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.

[15] 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 visit 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 use this dummy variable to identify individuals residing more than a 300-mile one-way driving distance from CAHA.

[16] Secondary factors for the 2005–2007 data choice are the timing to avoid potential confounds related to the Great Recession, and the impact of cell phones on sample representativeness. The latter point is

supported by the fact that MRIP is moving from phone to mail surveys in late 2017 due to the impact cell phones have had on sample representativeness and response rates.

[17] The intercept data contains a variable for the number of people per fishing party. The average is 2.73.

[18] For the phone survey, there were 124,255 individuals contacted in 2005, 148,206 in 2006, and 141,719 in 2007. These were phone interview contacts for people living in coastal counties from Delaware to Florida, and a majority did not take any trips.

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

[20] 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.

[21] This estimate is constructed using the formula $-1/\beta$ for value of a 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.

[22] The management plan and closure scenarios for *Alternative D* are provided in table A.1 in the online-only appendix.

[23] We also estimate models using data from 2009 and 2010 for robustness and find that the results are relatively similar to our preferred models using 2005–2007 data. These results are reported in the online-only appendix tables A.2 and A.3.

[24] 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}.

[25] Here we abstract from the distributional consequences of the policy.

[26] All NOAA MRIP data can be accessed here: http://www.st.nmfs.noaa.gov/recreational-fisheries/data-and-documentation/queries/index.