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EFFECT OF FORESTLAND AVAILABILITY BY OWNERSHIP TYPE ON LICENSE SALES FOR HUNTING: A SPATIAL APPROACH

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ABSTRACT. The effect of forestland availability under different ownership types on license sales for hunting in nine Southeastern states is empirically evaluated. An equation that represents license sales for hunting is estimated assuming the sale of hunting licenses in a particular county is related to the characteristics of that county as well as the characteristics and license sales for hunting in its neighboring counties. The positive effects of the amounts of both national and private forestland on license sales reaffirm the potential benefits of maintaining forestland to stimulate hunting. The positive

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spillover effect of national forests on license sales for hunting suggests that availability and close access to hunting in national forests within neighboring counties are important in supporting hunting license sales in a county. This study contributes to the general understanding of the drivers affecting individuals' decisions to use natural resources for hunting. Advances in natural resource modeling, specifically the spatial process model and geospatial data used in this research, make it possible to examine the interactions between the spatial dynamics and ownership attributes of the natural system, allowing policy makers to design natural resource management practices that respond to a system characterized by these interactions.

KEY WORDS: Hunting demand, hunting license sales, private forest, public forest, Southeastern United States, spatial interaction.

1. Introduction. Hunting is a popular cultural activity with a high economic impact in many parts of the world. The economic importance of hunting in the developed world has been reported in recent years (e.g., Grado et al. [2001], International Association of Fish and Wildlife Agencies [2002], Silberman [2002], Derek Murray Consulting Associates [2006], Hussain et al. [2007], Southwick Associates, Inc. [2007], Sharp and Wollscheid [2009], Munn et al. [2010], Bunnefeld et al. [2011]). Sharp and Wollscheid [2009] summarized the numbers of hunters and their expenditures in major developed regions. The authors reported that the numbers of hunters in the United States, Canada, Europe, and Australia are respectively 13.0, 1.2, 7.0, and 1.0 million and they are estimated to spend \$32.76 billion in total (i.e., 20.60, 0.72, 10.66, and 0.78 billion dollars in the United States, Canada, Europe, and Australia, respectively).[1]

Despite the importance of hunting in regional and local economies, hunting throughout most parts of the developed world has been dwindling for many years if not decades (Sharp and Wollscheid [2009]). For example, McCulloch et al. [1992] investigated the pattern of migratory bird hunting in Europe and found that the number of hunters of the majority of species had been in decline for over 40 years since 1950. Likewise, the US Fish and Wildlife Service [2007] reported that the hunting population in the Southeastern United States declined by 4%—from 4.9 to 4.7 million during the 1996–2006 period—a downward trend expected to continue in the future. Bowker et al. [1999] estimated that the US hunting population would decline by 11% during the 1995–2050 period, with hunting trips falling 22% in the South, a trend driven primarily by reduced interest in nature-based outdoor recreation among young adults (Cordell et al. [2008]). Specifically, young adults in the 21st century are not as engaged in nature-based activities as were baby boomers born during the demographic birth boom between 1946 and 1964 (Census News Release [2006]). Another reason for the decline is lack of access to public hunting lands (Mehmood et al. [2003]). Admittance to public hunting lands has declined because of an increasingly urban living environment because hunting is often not allowed in urban areas for safety reasons, and habitat destruction because urbanization has caused public hunting lands to decline (Poudyal et al. [2008b]). The decline may also be attributed to a cultural-detachment of growing urban populations from rural hunting traditions (Stedman and Heberlein [2001]).

The issue of better public access for hunting is not new and has been explored in various forms (Congressional Sportsmen's Foundation and Wildlife Management Institute [2002]). For instance, Poudyal et al. [2008a] suggested that the accessibility of public land has a significantly greater effect on license sales for hunting compared to the impact of its counterpart, private land. Specifically, a belief among the hunting community is that access to federal public land has been reduced through, e.g., infrastructure changes, road closures, lack of access across private land, area restrictions, and policy changes (CSF and WMI [2002]). Although Mozumder et al. [2007] suggested that the reduced per capita accessibility of public recreation land will expand the potential of lease-fee systems for recreational hunting on private land, other studies found that more land has been limited from public access for hunting than ever before (Duda et al. [2004], Jagnow et al. [2006]). Despite the mixed results, these studies emphasize the vital role of land ownership in examining the influence of public access to forestlands for hunting.

The objective of this research is to investigate the effects of hunting land availability by ownership type on hunting license sales. It is hypothesized that the accessibility of forestland under various ownership types (i.e., national, other federal, state and local, and private ownership types) has a critical effect on license sales. In this study, an equation modeling license sales for hunting is specified, where the quantity of resident hunting licenses sold is determined by fees, demographics, and other variables related with the characteristics and availability of hunting resources, including the availability of forestland under different ownership types. The equation is estimated with county-level data for resident hunting license sales, assuming the availability of hunting acreage in each county is represented by county forestland by ownership type.

2. Empirical model. Typically, an equation that represents license sales for hunting within a jurisdictional boundary is estimated to acquire the effects of fees, demographics, and other variables related with the characteristics and accessibility of hunting resources within that jurisdiction (e.g., Anderson et al. [1985], Brown and Connelly [1994], Sun et al. [2005]). An implicit assumption underlying that particular model specification is that license sales within a county are explained primarily by those variables within that county. This assumption may be untenable because hunting land is composed of large patches and natural ecosystems transcending political borders. Poudyal et al. [2008a] relaxed the supposition that county license sales are explained by own-county characteristics alone by using an arbitrary 100-mile buffer instead of county borders. However, neighborhood effects (or spillover effects) between a county's license sales and the characteristics of neighboring counties (defined as those that are contiguous to own counties in this research), including the license sales in those counties, were not explicitly modeled.

Although the literature using an equation representing county license sales that accommodates neighborhood effects is limited, the body of literature using regression models in other applications that capture the neighborhood effects of both dependent and explanatory variables (referred to as "spatial process model") has been growing quickly. In such models, the neighborhood effects are modeled as a weighted average of nearby cross-sectional units implemented by a matrix identifying neighborhood connections (referred to as "spatial weight matrix"; Cho et al. [2010]). This type of spatial process model has been applied in regression models using geographic information system (GIS) data because

Whittle's [1954] pioneering work of a spatial lag process model. Since then, extended applications, such as Cliff and Ord [1973, 1981], Anselin [1988], Anselin and Florax [1995] and Kelejian and Prucha [2010], have played a crucial role in the popularity of the spatial process model by incorporating a spatially lagged endogenous variable and a spatially autoregressive disturbance to mitigate error caused by neighborhood effects.

Following the general framework of the spatial process model, license sales for hunting in a particular county (or own county) are hypothesized to be related to the county's own characteristics and license sales in neighboring counties. Although hunters often reside, purchase licenses, and/or hunt in different counties, the purchase of a hunting license in a county represents an individual's demand for hunting rights whether or not those rights are exercised within that particular county. The equation representing license sales for hunting within a county is specified as:

$$\mathbf{y} = \lambda\mathbf{W}\mathbf{y} + \mathbf{X}\beta_{\mathbf{1}} + \mathbf{W}\mathbf{X}\beta_{\mathbf{2}} + \mathbf{Z}\beta_{\mathbf{3}} + \mathbf{W}_S\mathbf{Z}\beta_{\mathbf{4}} + \varepsilon, \tag{1}$$

where $\mathbf{y}$ is a vector of hunting license sales per capita, $\mathbf{W}$ is a row-standardized n by n contiguity matrix (or spatial weight matrix) with diagonal elements of zero and off-diagonal elements of 1 for all counties that are contiguous to own counties but are located within a state, $\mathbf{Wy}$ is a spatially lagged dependent variable that represents license sales in each county's neighboring counties, λ is a spatially lagged regressive term, $\mathbf{X}$ is a matrix of continuous and discrete variables that explain license sales excluding state dummy variables, $\mathbf{WX}$ is a spatially lagged matrix of continuous and discrete variables that represent the characteristics of each county's neighboring counties excluding state dummy variables, $\mathbf{Z}$ is a matrix of state dummy variables that accommodate state fixed effects on hunting license sales, $\mathbf{W}_S$ is a row-standardized n by n contiguity matrix (or spatial weight matrix) with diagonal elements of zero and off-diagonal elements of 1 for all counties that are contiguous to own counties and are located across the state border, $\mathbf{W}_S\mathbf{Z}$ is a spatially lagged matrix of state dummy variables that captures the effects in adjacent counties just across the state border on license sales for hunting, $\beta_{\mathbf{1}}$, $\beta_{\mathbf{2}}$, $\beta_{\mathbf{3}}$, and $\beta_{\mathbf{4}}$ are conformable coefficient vectors, and ε is a disturbance term.[2] A 5% level was chosen for significance level, and thus coefficients of variables are referred to as "significant" if the coefficients are statistically significant at the 5% level in the discussion of the empirical results.

Full information maximum likelihood (FIML) methods (Anselin [1988]) or general moment (GM) methods (Kelejian and Prucha [1999], Anselin and Lozano-Gracia [2008]) are usually used to estimate a regression model in the presence of spatially lagged variables. The GM method has a number of benefits over FIML (Kelejian and Prucha [2007]). First, the routine assumption of a normal distribution is relaxed. Second, the GM method avoids computation of an n by n determinant matrix, which is troublesome with larger data sets, even with advances in computing power. Despite these attractive features, when error terms are heteroscedastic, estimation with the GM method assuming homoscedasticity possibly will yield biased and inconsistent estimates of the parameters, which in turn may affect the main effect estimators of a model (Kelejian and Prucha [2010]).

The main steps of the GM method are comparable to the feasible generalized least squares method outlined by Kelejian and Prucha [1999], except heteroscedasticity of an unspecified form is assumed. The residuals from the GM method are tested for spatial error autocorrelation using a Lagrange multiplier (LM) test. The LM test is used frequently in the spatial econometric literature to detect spatial structure in the residuals of regression models (Anselin [1988]).

3. Data. In the United States, the economic impact of hunting is particularly high in the South, where a hunting tradition is strongly indicated by four of the top 10 US states with the most hunting activity being in the Southeastern regions of the United States (International Association of Fish and Wildlife Agencies [2005], Poudyal et al. [2008b]). Hunting is regarded as an intrinsic part of the Southern identity because it defines and is defined by the complex cultural, social, and technological forces of the South (Guignard [2004]). Munn et al. [2010] estimated the total economic impact of hunting in 13 Southeastern United States to be \$11 billion in output and 65,000 jobs created.[3] The total impact included a direct impact of \$7.6 billion in output and creation of 42,000 full or part time jobs, an indirect impact of \$1.8 billion in output and 8600 jobs, and an induced impact of \$1.6 billion in output and 14,500 jobs (Poudyal et al. [2007]).

The analysis included nine of the Southeastern states (Alabama, Arkansas, Georgia, Kentucky, Louisiana, South Carolina, Tennessee, Texas, and Virginia).[4] These states have 967 counties. After removing observations with missing data, we included 949 counties ($n = 949$) in the estimation of equation (1). State dummy variables were included to capture fixed, unobserved, state-specific factors using Tennessee as the reference.

We use three primary data sets: data for sales of hunting and fishing licenses, demographic data, and county-characteristics data including forest area (Table 1). We obtained county-level data for hunting license sales in 2000 from the state agencies accountable for recording these sales data, e.g., the Tennessee Wildlife Resource Agency. All categories of resident hunting licenses sold were summed and divided by total county population to obtain hunting license sales per capita. Similar data were acquired for per capita resident fishing license sales to serve as a complementary or substitute good for license sales for hunting. Combination hunting and/or fishing licenses were included in the sales data. Combination licenses were sold in all states; however, Tennessee was the only state requiring all hunters to purchase combination licenses (Bennear et al. [2005]).

A resident hunting license permits the holder to take legal game animals statewide by different hunting methods and resident fishing licenses permits the holder to take fish statewide by various fishing methods (Kentucky Department of Fish and Wildlife Resources [2010]). Qualifications for the purchase of resident hunting and/or fishing licenses differ by state but typically persons who possess a valid driver's license and have lived in the state for a minimum time period are eligible (Commonwealth of Kentucky [2008]).

The demographic data, including per capita income, employment status, age, race, and household type, were acquired from the 2000 US Census (U.S. Census Bureau [2000]). Per capita income was included to capture the effects of the opportunity cost of time for different income levels. The percentage of population holding full time jobs reflecting employment status was added to acquire the impact of economic status. Population percentages by age group (i.e., percentages of populations between 16 and 34 years of age and between 35 and 64 years of age) were included to provide the effects of different age cohorts. In

TABLE 1. Variable definition and descriptive statistics.

Variable	Description	Own county mean (standard error)	Neighboring counties mean (standard error)
Dependent variable			
Licenses sold per capita	Sum of all types of resident hunting permits sold divided by total population in 2000	0.12 (0.10)	
Demographic variables			
Personal income ($\times 10^{-3}$)	Per capita income	16.44 (3.45)	16.45 (2.55)
Employment	Percentage of population holding full time jobs (16+ years of age)	0.93 (0.02)	0.94 (0.02)
Age 16−34	Percentage of population between 16 and 34 years of age	0.25 (0.04)	0.25 (0.02)
Age 35−64	Percentage of population between 35 and 64 years of age	0.39 (0.03)	0.39 (0.02)
Caucasian	Percentage of population Caucasian	0.78 (0.17)	0.78 (0.14)
African American	Percentage of population of African American	0.16 (0.17)	0.16 (0.15)
Asian	Percentage of population Asian	0.01 (0.01)	0.01 (0.004)

(*Continued*)

TABLE 1. (*Continued*)

Variable	Description	Own county mean (standard error)	Neighboring counties mean (standard error)
Single-female headed households	Percentage of female-headed households (no husband present)	0.17 (0.06)	0.17 (0.05)
Single-male headed households	Percentage of male-headed households (no wife present)	0.05 (0.01)	0.05 (0.01)
Two-parent households	Percentage of parental households (reference)	0.78 (0.06)	0.78 (0.06)
Metro and gun-club variables			
Metro	1 if county is designated as metropolitan county, 0 otherwise	0.38 (0.49)	0.38 (0.28)
Gun clubs	1 if county has a gun club, 0 otherwise	0.03 (0.16)	0.03 (0.07)
Potential complement or substitute good variables			
Recreation	1 if county is designated as recreational county, 0 otherwise	0.04 (0.21)	0.04 (0.10)
Water-forest ratio	Water area divided by forest area	0.37 (4.61)	0.33 (3.68)
Fishing license	Total fishing permits sold divided by total population in 2000	0.15 (0.11)	0.15 (0.06)
Golf course ($\times 10^{-3}$)	Golf courses per capita	0.01 (0.02)	0.01 (0.01)

(*Continued*)

TABLE 1. (*Continued*)

Variable	Description	Own county mean (standard error)	Neighboring counties mean (standard error)
Amusement	1 if county contains alternative outdoor amusement, entertainment, or sport attractions, 0 otherwise	0.16 (1.13)	0.16 (0.50)
State dummy variables			
AL	1 if county is in Alabama, 0 otherwise	0.07 (0.26)	0.02 (0.15)
AR	1 if county is in Arkansas, 0 otherwise	0.08 (0.27)	0.01 (0.12)
GA	1 if county is in Georgia, 0 otherwise	0.17 (0.37)	0.03 (0.16)
KY	1 if county is in Kentucky, 0 otherwise	0.13 (0.33)	0.02 (0.14)
LA	1 if county is in Louisiana, 0 otherwise	0.07 (0.25)	0.16 (0.12)
SC	1 if county is in South Carolina, 0 otherwise	0.05 (0.21)	0.01 (0.11)
TN	1 if county is in Tennessee, 0 otherwise (Reference)	0.10 (0.30)	0.03 (0.18)
TX	1 if county is in Texas, 0 otherwise	0.25 (0.43)	0.01 (0.09)
VA	1 if county is in Virginia, 0 otherwise	0.10 (0.30)	0.01 (0.10)

(*Continued*)

TABLE 1. (*Continued*)

Variable	Description	Own county mean (standard error)	Neighboring counties mean (standard error)
Forest variables			
National forest	Per capita national forest area in acres	0.54 (2.60)	0.30 (0.85)
Federal forest	Per capita federal forest area other than national forest in acres	0.22 (0.86)	0.16 (0.33)
State and local government forest	Per capita state and local government forest area in acres	1.32 (24.91)	0.28 (0.98)
Private forest	Per capita private forest area in acres	17.25 (67.35)	7.42 (12.90)

addition, population percentages by three major races (i.e., Caucasian, African American, and Asian[5]) and population percentages by three major household types (i.e., single-female headed households, single-male headed households, and two-parent households) were respectively included to capture the effects of racial make up and different household types.

County-characteristics data including percentage of water and county typology, i.e., metro and recreational counties, were acquired from the USDA Economic Research Service (ERS [2009]). A dummy variable identifying metro and nonmetro counties was created based on the 2003 US Office of Management and Budget (OMB), which uses 2000 U.S. Census data (ERS [2007]). Metro counties were defined to consist of at least one urbanized area with a population of 50,000 or more plus adjacent territory having a high degree of social and economic integration with the core as measured by commuting ties (ERS [2007]). A dummy variable for recreational counties was created based on the recreation county typology developed by ERS (Johnson and Beale [2002]). Recreational counties were identified by ERS based on a weighted index converted from z-scores of three variables measuring employment, income, and seasonal housing to reflect recreational activity (See ERS [2007] to find out the details of how recreational counties were identified.).

Data on forest areas under different ownership types were obtained from the Forest Inventory and Analysis National Program (the "FIA program"; Bechtold and Patterson [2005]). The FIA program collects, analyzes, and reports information on relevant status and trends, including ownership classes of America's forests (Forest Inventory & Analysis [2005]). The ownership classes are grouped into national forest, other federal, state and local government, and private. National forest areas consist of the 9 million acres of federally owned national forests in the Southeastern states included in this study. Other federal forestlands include federal forestlands primarily owned by National Park Service, US Fish and Wildlife Service, the Departments of Energy and Defense, and the Tennessee Valley Authority. State and local government-owned forest areas are owned by state, county, and municipal local governments. Private forest is forestland owned by private individuals or organizations.

Forest areas by ownership class were collected separately for each state from FIA data. Two of the nine states collected the FIA data twice, i.e., Arkansas in 1995 and 2005 and Louisiana in 1991 and 2005. The data for the other states were collected by individual states over varying periods of time: Alabama (2000), Georgia (1998–2004), Kentucky (2000–2004), South Carolina (1999–2001), Tennessee (2000–2004), East Texas (2001–2003), West Texas (2007), and Virginia (1998–2001). The timing of hunting license sales data (2000) and data on forest areas did not exactly match except for Alabama. Temporal interpolation was used to approximate forest areas for the two periods in Arkansas and Louisiana. The forest areas in the six remaining states were used as proxies for the data in 2000.

Information about amusement parks, golf courses, and gun clubs was obtained from the National Outdoor Recreation Supply Information System (NORSIS [1997]). The NORSIS is primarily a secondary source county-level data system (Cordell and Betz [1997]). Data for amusement parks, gun clubs, and golf courses in 1997 were chosen to capture time-lagged effects on per capita license sales for hunting in 2000.

A dummy variable representing whether a county had one or more gun clubs was included to investigate whether gun-club affiliations positively affect license sales for hunting. The underlying hypothesis for its inclusion was that gun clubs allow easier participation in organized hunting trips and greater access to learning about opportunities for hunting. For this reason, counties with a gun club are hypothesized to have greater license sales for hunting than counties without one (Savannah Lakes Village [2011]). The gun-club effect was represented by a dummy variable instead of a continuous variable because the NORSIS data set indicated that the majority of the 949 counties in the study area had either one gun club (26 counties) or none, and only two counties had more than one gun club.

A dummy variable representing whether a county has additional outdoor leisure, entertainment, or sport attractions was used to capture potential substitute activities for nature-based outdoor activity such as hunting, particularly in urban areas (Gum and Martin [1975], Riess [1991], Poudyal et al. [2007]). The NORSIS category of amusement/entertainment/sports was used to create a dummy variable reflecting whether a county includes visitor attractions, primarily commercial in nature, that are not resource-based, historical, cultural, civic, or public, e.g., memorials and monuments, or shopping outlet malls and the like.

4. Empirical results. The estimates of the per capita hunting license sales model are presented in Table 2. The null hypothesis of no spatial autocorrelation in the error term is not rejected (LM $= 0.268$; p value > 0.1), signifying that the error terms are unlikely to be inefficient. The spatially lagged dependent variable representing license sales in neighboring counties (λ) is statistically significant at the 5% level. This result suggests that license sales in a given county are concurrently affected by license sales in neighboring counties (spillover effect).

The significant coefficient for the percentage of Caucasian population indicates that an increase in the Caucasian population by 1% increases own-county license sales for hunting by 1.66 licenses/1000 persons. The positive effect of Caucasian populations is a frequent

TABLE 2. Model estimation results for the demand for hunting license sales.

Variables	Marginal effect of own county (standard error)	Marginal effect of neighboring counties (standard error)
Demographic variables		
Personal income ($\times 10^{-3}$)	−0.001	−0.00005
	(0.001)	(0.002)
Employment	0.001	−0.0005
	(0.002)	(0.003)
Age 16−34	−0.161	0.061
	(0.097)	(0.231)
Age 35−64	−0.140	−0.119
	(0.165)	(0.331)
Caucasian	0.166[a]	−0.197
	(0.079)	(0.129)
African American	0.089	0.191
	(0.084)	(0.131)
Single-female headed households	−0.330	−1.272
	(0.338)	(0.682)
Single-male headed households	0.085	0.089
	(0.100)	(0.205)
Two-parent households	0.163	−0.916
	(0.200)	(0.480)
Metro and gun-club variables		
Metro	−0.017[a]	0.0005
	(0.006)	(0.012)
Gun clubs	−0.006	−0.039
	(0.012)	(0.034)

(*Continued*)

TABLE 2. (*Continued*)

Variables	Marginal effect of own county (standard error)	Marginal effect of neighboring counties (standard error)
Potential complement or substitute good variables		
Recreation	−0.025[a]	0.024
	(0.011)	(0.026)
Water-forest ratio	0.001	−0.003
	(0.001)	(0.002)
Fishing license	0.224[a]	−0.095
	(0.025)	(0.059)
Golf course ($\times 10^{-3}$)	−0.191	1.438[a]
	(0.124)	(0.353)
Amusement	−0.002	−0.007
	(0.002)	(0.005)
State dummy variables		
AL	−0.038[a]	−0.017
	(0.015)	(0.015)
AR	−0.001	0.003
	(0.015)	(0.020)
GA	−0.057[a]	−0.004
	(0.013)	(0.015)
KY	0.079[a]	0.021
	(0.013)	(0.016)
LA	0.104[a]	0.001
	(0.017)	(0.019)
SC	−0.042[a]	−0.012
	(0.017)	(0.020)
TX	−0.010	−0.005
	(0.016)	(0.024)

(*Continued*)

TABLE 2. (*Continued*)

Variables	Marginal effect of own county (standard error)	Marginal effect of neighboring counties (standard error)
VA	0.111[a]	−0.021
	(0.015)	(0.022)
Forest variables		
National forest ($\times 10^{-2}$)	0.798[a]	0.910[a]
	(0.092)	(0.337)
Other federal forest ($\times 10^{-2}$)	−0.362	−0.048
	(0.250)	(0.735)
State and local government forest ($\times 10^{-2}$)	−0.010	0.214
	(0.013)	(0.442)
Private forest ($\times 10^{-2}$)	0.017[a]	−0.017
	(0.005)	(0.038)
Spatial variable		
Lambda (λ)	0.241[a]	
	(0.030)	

[a] Indicates statistical significance at the 5% level.

occurrence in hunting license sales models (Poudyal et al. [2008a]). Often, Caucasian populations are found to have higher demand for hunting trips (Mehmood et al. [2003]). Metro counties have 1.7% lower license sales per capita than nonmetro counties. This finding suggests that city dwellers are less likely to be engaged in hunting than rural residents. Recreational counties have 2.5% lower license sales per capita than nonrecreation counties. This result implies that recreational attractions serve as substitutes for hunting activities. Rather than being substitutes, hunting and fishing are complements in the own-to-own county relationship. A rise in fishing license sales in a county by 1 licenses/1000 persons increases license sales for hunting by

0.224 licenses/1000 persons in the same county. Although the data include combination hunting/fishing licenses, in all states but Tennessee, purchase of these licenses is voluntary. (The mandatory purchase of combination licenses in Tennessee and other differences among states are captured by the state dummy variables.)

The state dummy variables show that Alabama, Georgia, and South Carolina have statistically lower hunting license sales per capita than Tennessee, whereas Kentucky, Louisiana, and Virginia have higher license sales. Holding all else constant, license sales per capita in Alabama, Georgia, and South Carolina are 38, 57, and 42 licenses/1000 persons fewer than those in Tennessee, whereas license sales in Kentucky, Louisiana, and Virginia are 79, 104, and 11 licenses/1000 persons greater than in Tennessee.

A decrease in own-county national forestland by 10 acre/1000 persons decreases own-county license sales for hunting by 7.98 licenses/1000 persons whereas the same decrease in neighboring counties decreases own-county license sales by 9.10 licenses/1000 persons. Assuming county mean hunting license sales of 120/1000 persons and mean national forestland of 540 acres/1000 persons in the own county, a decrease of national forestland in the own county by 1% (or 5.4 acres/1000 persons) would decrease own-county license sales by 3.6% (or 4.3 hunting licenses/1000 persons). Assuming the same mean license sales and mean national forestlands of 300 acres/1000 persons in neighboring counties, a decrease of national forestland in neighboring counties by 1% (or 3 acres/1000 persons) would decrease own-county license sales by 2.3% (or 2.73 hunting licenses/1000 persons). A decrease in own-county private forestland by 10 acres/1000 person decreases own-county license sales for hunting by 0.17 licenses/1000 persons. Contrary to the result for national forestland, private forestland in neighboring counties does not affect own-county license sales per capita for hunting.

5. Discussion. The effect of forestland availability under different ownership types on license sales for hunting in nine Southeastern states is empirically evaluated. An equation that represents license sales for hunting is estimated assuming the sale of hunting licenses in a particular county is related to the characteristics of that county as well as the characteristics and license sales for hunting in its neighboring counties. The positive effects of the amounts of both national and private forestland on license sales reaffirm the potential benefits of maintaining forestland to stimulate hunting. The positive spillover effect of national forests on license sales for hunting suggests that availability and close access to hunting in national forests within neighboring counties are important in supporting hunting license sales in a county.

The smaller marginal effect of private forestland (or the additive change in license sales per capita for hunting for a unit change in private forestland) compared to national forestland may result from more limited access to private forests because of lease fees and legal barriers to hunting on those lands. This result confirms previous findings that the accessibility of public land has a significantly greater effect on license sales per capita for hunting compared with the impact of its counterpart, private land (Poudyal et al. [2008a]). In addition, the owners of private lands in many states are not required to have licenses to hunt game (e.g., pheasants) on their own properties (e.g., Alabama, Georgia, and Tennessee), which may decrease the marginal effect of private forestland.

The insignificant effects of other federal government forestlands on both own-county and neighboring county license sales per capita for hunting are not surprising because hunting is not allowed on most federal lands primarily owned by National Park Service, US Fish and Wildlife Service, the Departments of Energy and Defense, and the Tennessee Valley Authority (e.g., Public Hunting Lands [2010]). The insignificance of state and local government forestland is mostly because of its unsuitability as hunting ground because forestlands owned by state and local governments are usually a relatively small proportion of all forestland.

Hunting license among states is not comparable in general because hunting regulations are structured differently among states. For example, Georgia has 20 nonpermanent resident licenses (e.g., alligator hunting, big game annual, dog deer hunting) whereas Tennessee has 16 nonpermanent resident licenses (e.g., waterfowl supplemental, big game gun supplemental). Relatively higher license sales per capita for hunting in Kentucky, Louisiana, and Virginia may be related with abundant hunting opportunity for game species. For example, Louisiana offers attractive hunting sites by providing large size bucks with good antlers because of limited hunting pressure in conjunction with the swamp-laced farmland along the Mississippi River and its major tributaries (North American Whitetail Magazine [2004]).[6]

At least two interesting points merit further discussion. First, license sales for hunting in a particular county are affected by that county's characteristics as well as the license sales and characteristics of its neighboring counties. This model is designed to capture the neighborhood effects of adjacent counties as well as own-county characteristics. Neighborhood effects are difficult to identify because hunting ground is composed of large patches and natural ecosystems transcending political borders, and many hunters travel long distances across several counties to hunt. Future research could focus on developing models that identify the scope over which license sales for hunting are affected by neighborhood effects. For example, a spatial model that systematically identifies distances to hunting grounds from hunters' locations of origin could help identify this range of significant neighborhood effects. Specifically, travel time that uses road network data, including information about travel impediments (i.e., speed limits), impactors (i.e., stop signs and traffic lights), and landscapes (i.e., elevations) may provide a better measure of distance than the straight-line measure.

The second point of interest concerns the difference in neighborhood effects between national and private forestlands. The lack of interaction between private forestland in neighboring counties and own-county license sales for hunting reveals an opportunity for private forestland owners to stimulate local license sales for hunting. Specifically, a local cooperative structure that enhances the cross-county coordination among private forestland owners could improve license sales by advancing the interaction of private forestland holdings across counties. For example, lease-fee systems for recreational hunting on private land could be jointly structured to foster a cooperative approach among neighboring landowners. Burden [2009] suggests that a local cooperative structure that fosters coordination among neighboring landowners, working toward common wildlife management goals, could improve license sales for hunting. In some regions, interest in hunting may be sufficient to stimulate such coordination among private forestland owners. For example, license sales for hunting around the Atlanta metro area and counties in Southwest Georgia have been enhanced by Quality Deer Management regulations on mixed forestlands and soybean and corn agricultural lands, generating tremendous interest in and revenue from lease-fee hunting (Hamilton et al. [1995], Green and Stowe [2000]). Hence, private benefits from coordination of hunting on private lands in that area may be sufficient to encourage cooperation among landowners. For instance, coordination among neighboring forest owners across counties could benefit by offering a cooperative lease-fee system that provides a quality-controlled deer population within a larger hunting area.

The implication of this study for the general understanding of the drivers affecting individuals' decisions to use natural resources for hunting can be summarized as: The decision to use natural resources for any purpose involves a complex process driven by interactions between the spatial dynamics and ownership of the natural systems (Johnston et al. [2010], Bunnefeld et al. [2011], Fulton et al. [2011]). Management practices for natural resource use without consideration of these interactions do not respond to the system dynamics caused by the interactions Walters and Hilborn [1976], Packer et al. [2009], (Fulton et al. [2011]). Advances in natural resource modeling, specifically the spatial process model and geospatial data used in this research, make it possible to examine the aforementioned interactions, allowing policy makers to design natural resource management practices that respond to a system characterized by these interactions. Our suggestion of a local cooperative structure to enhance cross-county coordination among private forestland owners to stimulate hunting is an example of an inference drawn from the spatial modeling of hunting license sales in this research. Although improving the apparent realism of natural resource modeling alone may not necessarily capture all factors driving consumer decisions about natural resource use (Walters [1985]), this research emphasizes that such efforts are important for developing more effective natural resource management practices in the future.

ENDNOTES

1. Twenty-five EU member states and 10 countries where Federation of Associations for Hunting and Conservation of the EU (FACE) has members.

2. The equation (1) resembles a "spatial Durbin" type model. See Pace and LeSage [2010] and Pace et al. [2011] for the methodological issues associated with this spatial Durbin type model.

3. The 13 Southeastern states in the study include: Alabama, Arkansas, Florida, Georgia, Kentucky, Louisiana, Mississippi, North Carolina, Oklahoma, South Carolina, Tennessee, Texas, and Virginia.

4. Four of the 13 southeastern states, i.e., Florida, Mississippi, North Carolina, and Oklahoma, were excluded from the analysis because of missing observations.

5. Hispanics was introduced in the 1980 census as a category of ethnicity, separate and independent of race. Hispanics make up a racially diverse group, which includes Caucasian and African American.

6. Due to lack of adequate variability in hunting license fees across observations (because the license fee is fixed at the state level), license fees were not included in the model. Instead, the state dummy variables capture differences in license fees as well as other factors mentioned here that affect hunting license sales across the states.

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