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The Dynamic Effects of Open-Space Conservation Policies on Residential Development Density

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The Dynamic Effects of Open-Space Conservation Policies on Residential Development Density

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Abstract: Recent economic analyses emphasize that designated open-space increases the rents on neighboring residential land, and likewise, the probability of undeveloped land converting to residential uses. This paper addresses a different question: What is the effect of local open space conservation on the rate of growth in the density of existing residential land? The analysis is relevant for exurban development and also for remote lakeshore development, where shoreline development density can rapidly increase over time and open-space policies are often advocated as a way to protect ecosystems by reducing development. A discrete choice econometric model of lakeshore development is estimated with a unique parcel-level spatial-temporal dataset, using maximum simulated likelihood to account for i) the panel structure of the data, ii) unobserved spatial heterogeneity, and iii) sample selection resulting from correlated unobservables. Results indicate that, contrary to the intuition derived from the current literature, local open space conservation policies *do not increase* the rate of growth in residential development density, and some open space conservation policies may *reduce* the rate of growth in residential development density. This is consistent with land-value complementarity between local open space and parcel size. Spatially-explicit simulations at the landscape scale examine the relative effects of conservation policies on the time path of development.

Keywords: spatial modeling, land-use change, open space, development density, shoreline development, zoning.

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The Dynamic Effects of Open-Space Conservation Policies on Residential Development Density

1. Introduction

Recent studies of land development have examined the question of how open-space conservation efforts such as conservation easements affect the conversion of agricultural and forest land to residential development (Wu and Plantinga 2003; Turner 2005; Armsworth et al. 2006; Wu 2006). This literature emphasizes a point that land use planners and land conservation organizations often overlook: By making the local landscape more attractive, local open-space conservation may actually *increase* the rate of nearby land development. The underlying economic logic is that open space conservation increases the value of land in residential development, but has little or no effect on the value of land in agriculture or forestry, and so it effectively increases the probability that any particular agricultural or forestry parcel is converted to residential. These studies are supported by a number of econometric studies of land use conversion, including Bockstael (1996), Irwin and Bockstael (2004), and Walsh (2007).[1]

The existing literature typically treats land conversion as a binary process: agricultural or forest land converts to a fixed residential development density. But residential development often becomes increasingly dense over time, which leads to the question addressed in this paper: What is the effect of open space conservation policies on the rate of change in residential density? Drawing on the existing economics literature on land conversion, one might reasonably conclude that such policies stimulate higher residential development densities. In this paper we argue that this is not necessarily the case, and we apply a parcel-level econometric model to a unique spatial-temporal dataset to show that in at least one instance –shoreline development in northern Wisconsin –both open space in the form of public conservation land on shorelines, and maximum development density restrictions in the form of minimum shoreline frontage requirements, *reduce* the rate of residential subdivision.

Cast most generally, ambiguity about the effect of open space on residential development density arises because the factors such as local open space that increase the value of residential land affect *both* the returns from subdividing a residential parcel *and* the returns from keeping the parcel in its original state. Open space and other neighborhood attributes (local public goods) are weakly complementary with residency in the neighborhood –residency is essentially required for their consumption. This implies that a fixed premium attaches to every residential parcel in the neighborhood. It is the existence of this premium that explains why open space accelerates the conversion of agricultural and forest land to residential development. It also presents parcel owners with the opportunity to increase their welfare by creating *more* rather than fewer new parcels upon development. If the value of a parcel is separable in open space, or if parcel size and open space are substitutes in the land value function, then the parcel owner maximizes land value by subdividing to the fullest extent allowed by zoning and the natural features of the parcel. On the other hand, if the decision context is the further subdivision of a parcel that is already in residential use, and in the land value function the size of a parcel is complementary to open space, then it becomes possible that an increase in local open space serves to delay subdivision, and that the number of parcels created upon subdivision is lower than feasible under relevant zoning law and the natural features of the original parcel. The explanation is that the incentive to capture the "open space premium" associated with each new parcel is mitigated by the positive effect of open space on the marginal value of parcel size.

Understanding how open-space conservation affects the dynamics of residential development is important for its obvious implications for economic welfare and land use planning, and also for its implications for ecosystem change, perhaps especially in the case of lakeshore development (the empirical application examined in this study). The development of shoreline property can result in major ecosystem change across North American lakes. In particular, high density shoreline development can lead to the clearing of sunken logs serving as habitat for a variety of aquatic species (Christensen et al. 1996), reduced growth rates of fish (Schindler et al. 2000), reduced abundance of amphibians and birds (Woodford and Meyer 2003; Lindsay et al. 2002), increased nutrient loading of lakes (Schindler 2006), and an increase in aquatic species invasions arising from increased recreational use of lakes (Hrabik and Magnuson 1999).

In this paper we analyze the effects of shoreline zoning restrictions and public conservation land on the amount and spatial configuration of land development across a fast-growing lake system in the northern forest region of Wisconsin. The econometric model is estimated with an extensive panel dataset developed by reconstructing historical GIS data from paper plat maps. The development process of 1,575 privately-owned shoreline parcels is followed from 1974 through 1998 in four-year intervals, resulting in a unique spatial-temporal dataset on land development over a 25-year period for 140 individual lakes.

In addition to the unique spatial-temporal dataset, there are two distinguishing features of the econometric modeling that contribute to the land use literature. The first is the joint estimation of subdivision density –that is, the number of parcels created per unit shoreline –with the binary decision of whether or not to subdivide in the first place. Most studies analyzing the probability of residential development assume that development occurs at the maximum density allowable by zoning (Bockstael 1996; Irwin and Bockstael 2002; Carion-Flores and Irwin 2004).[2] With our dataset this presumption that subdivision occurs at maximum density is not justified: 50% of all observed subdivisions generated a lower density than allowed by law.[3] Given our theoretical framework, joint rather than separate estimation of these decisions is necessary because the two decisions embed correlated random variables. This presents a classic sample selection problem: the researcher observes the number of new parcels created upon subdivision only when subdivision actually occurs. The full-information maximum simulated likelihood approach used in this paper explicitly accounts for sample selection that arises from this decision problem.

The second distinguishing feature of the econometric modeling is the use of a random effects framework to account for both the panel structure of the data and potential unobserved spatial heterogeneity. The development decision depends on attributes unobservable to the analyst. One can expect that these unobservables are correlated over time and across parcels on the same lake. This implies that repeated observations of a landowner's development decision are temporally correlated at the parcel level, and that development decisions across parcels are correlated at the lake level. We develop a random effects model to account for such temporal and spatial correlation that can be estimated within the full-information maximum simulated likelihood framework discussed above.

The econometric analysis investigates the effect of two open space conservation policies on lakeshore development. The first policy is the creation of public conservation land, and the second is a zoning policy specifying the minimum shoreline frontage required for a new shoreline parcel (this is analogous to a minimum lot size requirement). Each of these open space policies provides the foundation for a pair of tests of whether parcel size and open space are separable, substitutes, or complements in the land value function. The first test concerns the effect of open space on the decision to subdivide, and the second test concerns the decision about the number of new lots created upon subdivision. The overall conclusion of the analysis is that because open space and parcel size are apparently complements in the land value function, open space conservation *does not increase* the rate of growth in residential development density, and open space conservation via the creation of public conservation land may actually *reduce* the rate of growth in residential development density.

The paper is organized as follows. A simple exposition of the landowner's subdivision problem is presented in section 2 to contrast the problem examined here with the usual subdivision problem examined in the literature. The econometric model and estimation framework are presented in section 3, and the application is described in section 4. Estimation results are presented in section 5, and a landscape simulation model is developed in section 6 to analyze the effects of zoning and publicly-owned shoreline on the spatial and temporal patterns of land conversion. The paper concludes in section 7 with a discussion of the implications of the analysis for policies designed to control land development.

2. Exposition of the landowner's subdivision problem

Drawing on Capozza and Helsley (1989), the recent literature on land development typically assumes that the rental value of undeveloped land is constant, and the value of developed land increases smoothly over time. As a result the development decision is a deterministic optimal stopping problem in which development takes place at time t when,[4]

$$R^{D}\left(w,t\right)=R^{UD}\left(z\right), \tag{1}$$

where $R^{D}\left(w,t\right)$ is the rental value of developed land, $R^{UD}\left(z\right)$ is the rental value of undeveloped land, w is a vector of variables affecting the value of developed land, and z is a vector of variables affecting the value of undeveloped land. Although the econometric decision model we develop below is more general than this –in particular, the decision problem is stochastic and the rental value of developed land depends on the number of new parcels created, and thus is itself the outcome of a maximization problem (how many parcels to create upon subdivision) –it is instructive to briefly compare such an optimal stopping problem with its counterpart applicable in the context of our analysis.

The standard assumption is that the rental value in the undeveloped state is not a function of the same set of variables as the rental value in the developed state $\left(w \neq z\right)$, because the undeveloped state is typically forestry or agriculture. The implication of this is that any variable $w_j \in w$ that increases the rental value of the developed state will *decrease* the time to land development (or, analogously in a stochastic setting, *increase* the probability that an undeveloped parcel is developed in the current period).

In the context of our analysis, where the decision choice involves converting a residential parcel into two or more residential parcels, the rental value in the "undeveloped" state (un-subdivided parcel) is a function of the same variables that affect the rental value in the developed state. Formally, we designate *f* as the variable over which subdivision occurs; *f* is lake frontage in our econometric application, and it is parcel area in most settings. The vector of determinants of rental value is then expanded to $w = \left(f, x\right)$, where *x* is a public good (like open space), the value of which therefore accrues to all parcels created upon subdivision. Assuming that conditional on development, the rental value of the land is maximized with just two parcels, the rental value in the developed state can be presented as $R_1 + R_2 = R\left(f_1, x, t\right) + R\left(f_2, x, t\right)$, and the rental value in the undeveloped state can be presented as $R_{Tot} = R\left(f_1 + f_2, x, t\right)$.

Now assuming that the value of the parcel in its "dense" state of development is rising faster than its value in its "sparse" state of development –an assumption we maintain to keep the analysis parallel to the Capozza and Helsley model –the condition for subdivision is now,

$$R_1 + R_2 = R_{Tot} \quad . \tag{2}$$

In the context of our problem it is no longer obvious that an increase in the public good will reduce the time to development. Differentiation of (2) with respect to x and t generates the result,

$$\left[\left(\frac{\partial R_1}{\partial x}+\frac{\partial R_2}{\partial x}\right)-\frac{\partial R_{Tot}}{\partial x}\right]dx=\left[\frac{\partial R_{Tot}}{\partial t}-\left(\frac{\partial R_1}{\partial t}+\frac{\partial R_2}{\partial t}\right)\right]dt \quad . \tag{3}$$

Staying with the Capozza and Helsely (1989) assumption that rents from the developed state are rising faster than rents from the undeveloped state, the bracketed term on the right hand side of (3) is negative. For the case where R is separable in x –the case where the public good fetches a simple premium for any parcel in the neighborhood –the bracketed term on the left-hand side of (3) is positive, and an increase in x must decrease the time to development (increase the probability of development). For the case where f and x are complements (that is, $\frac{\partial R^2(f,x,t)}{\partial f\,\partial x} > 0$), it becomes possible that the bracketed term on the left-hand side of (3) is actually negative, in which case an increase in x will *increase* the time to development (reduce the probability of development), because an increase in x increases the marginal value of frontage more on a larger lot.

The important insight from this simple analysis is that because a local public good like open space may be complementary to parcel size, an increase in the public good, though it increases the value of residential land, does not necessarily induce an increase in the likelihood that a parcel already in residential use is further subdivided. Analogous reasoning makes clear that an increase in the local public good also may not induce an increase in the number of parcels created upon subdivision; more public-good premiums are generated by creating more (and smaller) parcels, but possibly the value of each premium declines as more parcels are created.

3. Econometric model of the landowner's subdivision decision

We cast a lake shoreline owner's decision problem as a matter of deciding how many parcels to create at time t. We do not formally model the dynamics of this decision problem, instead casting the decision problem in terms of the (reduced form) net value of creating m_t new parcels at time t, $m_t = 1,2,\ldots,$ with the dynamics of the decision problem implicitly embedded in the land value function via the presence of important state variables, such as shoreline development density, as arguments of the land value function.

Parcel n on lake l is subdivided at time t if the net land value of subdivision is positive. Formally, we denote this land value by,

$$U\left(w_{nt}\right)+\mu_l\,, \tag{4}$$

where w_{nt} is a set of parcel characteristics (including characteristics of the lake on which the parcel sits, such as size of the lake), and μ_l denotes a lake-specific characteristic observed by the parcel owner but not by the analyst. We model μ_l as an iid normal random variable distributed with mean zero and standard deviation σ_1.

The land value function in (4) is itself an indirect function derived from the decision about how many new parcels to create, given subdivision occurs. Formally, the value of creating m new parcels from parcel n at time t is given by,

$$V_m\left(w_{nt}\right)+\mu_l+\varphi_{mnt}\,, \tag{5}$$

where φ_{mnt} is a decision-specific variable observed by the parcel owner at the time the decision is made, but a random variable from the perspective of the analyst. Given the decision to subdivide, the decision about the number of parcels to create is the solution to the problem,

$$U\left(w_{nt}\right) = \max\left\{V_m\left(w_{nt}\right) + \varphi_{mnt}\right\}_{m=1}^{M}, \tag{6}$$

where the land value function U is a random variable because it is derived by maximizing over a set of random variables. For instance, if φ_{mnt} has a Type I extreme value distribution with location parameter equal to zero and a common scale parameter ξ for $m=1,\ldots M$, then

$$U\left(w_{nt}, \upsilon_{nt}\right) = \frac{1}{\xi}\left(\ln\left[\sum\nolimits_M V_m\left(w_{nt}\right)\right] - \gamma\right) + \upsilon_{nt}, \tag{7}$$

where υ_{nt} is distributed Type I extreme value with location equal to zero and scale equal to ξ, and γ is Euler's constant. This suggests a rather complicated form for U that generally must be derived by simulation if one chooses to specify particular forms of $V_m\left(w_{nt}\right)$ and particular distributions of φ_{mnt}. Moreover, (6) and (7) make clear that explicit derivation of U requires parameterization of the functions $V_m\left(w_{nt}\right)$, which significantly increases the size of the econometric problem by adding parameters that are not of first-order importance to the analysis. With this in mind, we instead simply assert that (6) generates a land value function adequately represented by a function of the form,

$$U\left(w_{nt}, \upsilon_{nt}\right) = \delta w_{nt} + \upsilon_{nt}, \tag{8}$$

where υ_{nt} is an iid standard normal random variable.

The number of parcels created upon subdivision is defined by the function,

$$m^*\left(w_{nt}, \omega_n\right) = \arg\max_m \left\{V_m\left(w_{nt}\right) + \varphi_{mnt}\right\}_{m=1}^{M}, \tag{9}$$

where the variable ω_n is a random variable whose presence in $m^*(\cdot)$ reinforces that m^* is a random variable by virtue of the fact that it is generated by an operation on the set of random variables φ_{mnt}. Once again taking the tack that we can adequately represent this indirect function by a simple specification capturing the essential elements of the decision problem, we assume that m^* is Poisson-distributed, that its expected value depends on w_{nt} and the random variable ω_n, and that it is necessarily correlated with the land value of subdivision $U(w_{nt}, \upsilon_{nt})$ by virtue of the fact that both U and m^* are derived from operations on the same set of random variables φ_{mnt}. In particular, we specify that the expected value of m^* takes the zero-censored exponential form,

$$\begin{aligned} Em^* &= \exp(\theta w_{nt} + \omega_n) / \left[1 - \exp(-\exp(\theta w_{nt} + \omega_n)\right] \\ &= \exp(\theta w_{nt} + \sigma_2 \eta_n) / \left[1 - \exp(-\exp(\theta w_{nt} + \sigma_2 \eta_n)\right] \end{aligned}, \tag{10}$$

where ω_n is a normal random variable with standard deviation σ_2, and so it follows that η_n is a standard normal random variable. To account for the correlation of U and m^*, we assume that υ_{nt} and η_n are jointly normal:

$$\{\eta_n, \upsilon_{nt}\} \sim N\left[(0,0),(1,1,\rho)\right] \quad . \tag{11}$$

The probability that $m^* = m$, *m=1,2…,* given that subdivision occurs, and conditional on w_{nt} and η_n, is

$$\Pr\left[m^* = m \mid w_{nt}, \eta_n\right] = \frac{e^{-\exp(\theta w_{nt} + \sigma_2 \eta_n)} \left(\exp(\theta w_{nt} + \sigma_2 \eta_n)\right)^m}{m!(1 - \exp(\theta w_{nt} + \sigma_2 \eta_n))} \tag{12}$$

It deserves emphasis that ignoring in estimation the obvious correlation between U and m^* would generate inconsistent as well as inefficient estimators because of the censored nature of the data: we observe m^* only for those cases where subdivision takes place. The particular formulation used here –a probit model for the subdivision decision, and a Poisson model for the number of parcels created, with correlated errors across the models –can be estimated by applying the formulation used by Greene (2006) to address the sample selection issue implicit in such data.[5] As stated by Greene, if the random element of the binary selection model (decision to subdivide) is uncorrelated with the random element of the count model (number of parcels to create), there is no issue of selection bias. But as shown above, in our particular analysis such correlation is true by the construction of the underlying decision problem, and so it is essential to explicitly model the correlation in the unobservables of the models.

3.1. Estimation of the decision parameters

The decision model involves a number of parameters to be estimated from the data: $\delta, \theta, \sigma_1, \sigma_2$, and ρ. The data used in the analysis includes observations on the decision to subdivide, y_{nt}, where $y_{nt} = 1$ if the net value of subdivision defined in (8) is positive; property characteristics w_{nt}; and, given that subdivision takes place, the number of parcels created, m^*_{nt}. Our estimation approach extends the selection framework developed by Greene (2006) to include a random effects structure. Letting $\Phi(\cdot)$ denote the standard normal cumulative distribution function, the probability of subdivision conditional on w_{nt} and μ_l is given by

$$\Pr\left(y_{nt} = 1 \mid w_{nt}, \mu_l\right) = \Phi\left(\delta w_{nt} + \mu_l\right), \tag{13}$$

and so given the properties of a joint normal distribution, the probability of subdivision conditional on w_{nt}, μ_l, and η_n is given by

$$\Pr\left(y_{nt} = 1 \mid w_{nt}, \mu_l, \eta_n\right) = \Phi\left(\left[\delta w_{nt} + \mu_l + \rho\eta_n\right] \big/ \sqrt{1-\rho^2}\right). \tag{14}$$

Conditional on η_n and μ_l, the decision to subdivide and the number of parcels created upon subdivision are statistically independent, and so the probability of observing a subdivision that creates m parcels is simply the product of (12) and (14). Conditioning this probability on only the observed variables w_{nt} requires integrating out η_n and μ_l :

$$\Pr(y_{nt}=1, m_{nt} \mid w_{nt}) = \iint \left[\frac{e^{-\exp(\theta w_{nt}+\sigma_2\eta_n)} \left(\exp(\theta w_{nt}+\sigma_2\eta_n)\right)^m}{m!(1-\exp(\theta w_{nt}+\sigma_2\eta_n))} \right] \cdot \left[\Phi\left(\left[\delta w_{nt}+\mu_l+\rho\eta_n\right] \big/ \sqrt{1-\rho^2} \right) \right] \phi(\eta_n)\phi(\mu_l) d\eta_n d\mu_l \quad ,(15)$$

where $\phi(\eta)$ and $\phi(\mu)$ are the density functions for η and μ . The probability of the observed behavior on parcel n at time t is generally stated,

$$\Pr(y_{nt}, m_{nt} \mid w_{nt}) = \iint \left[(1-y_{nt}) + y_{nt} \cdot \frac{e^{-\exp(\theta w_{nt}+\sigma_2\eta_n)} \left(\exp(\theta w_{nt}+\sigma_2\eta_n)\right)^m}{m!(1-\exp(\theta w_{nt}+\sigma_2\eta_n))} \right] \cdot \left[\Phi\left((2y_{nt}-1) \cdot \left[\delta w_{nt}+\mu_l+\rho\eta_n\right] \big/ \sqrt{1-\rho^2} \right) \right] \phi(\eta_n)\phi(\mu_l) d\eta_n d\mu_l , \quad (16)$$

where the term $2y_{nt}-1$ in the density function of the standard normal is an expositional and computational convenience that exploits the symmetry of the normal distribution.

At this juncture it is important to emphasize that observations of the subdivision decision are not statistically independent because we have included two random variables that capture unobservable effects that persist across observations. In particular, η_n captures parcel-level unobservables that persist over time, and μ_l captures lake-level unobservables that persist over time and across all parcels on the same lake. Inclusion of these variables compels simulation of the likelihood function because direct calculation would involve multi-dimensional integration (1+ the number of parcels on lake l) over an inevitably tiny probability space.

We denote by N_l the set of sample properties on lake l, and we denote by D_l the full set of subdivision decisions (m, y) made by members of N_l over all time periods. Conditional on η_n and μ_l, the probability of the observed subdivision decision on parcel n at time t is simply the integrand of (16), and so, conditional on η_n and μ_l , the probability of D_l is,

$$\Pr(D_l)=\prod_{n\in N_l}\prod_t\left[(1-y_{nt})+y_{nt}\cdot\frac{e^{-\exp(\theta w_{nt}+\sigma_2\eta_n)}\left(\exp(\theta w_{nt}+\sigma_2\eta_n)\right)^m}{m!(1-\exp(\theta w_{nt}+\sigma_2\eta_n))}\right]\cdot\left[\Phi\left((2y_{nt}-1)\cdot\left[\delta w_{nt}+\mu_l+\rho\eta_n\right]\big/\sqrt{1-\rho^2}\right)\right]\cdot \tag{17}$$

The likelihood of the observed subdivision behavior on lake l can be simulated by drawing randomly from the independent normal distributions of η and μ. Taking R sets of draws, with each set comprised of a single draw from the distribution of μ and N_l draws from the distribution of η, generates an approximation of the likelihood function,

$$\Pr^{Sim}\left(D_l\right)=\frac{1}{R}\Pr\left(D_l\right)\ . \tag{18}$$

The full simulated log likelihood function is then $\sum_{l=1}^{L}\log[\Pr^{Sim}(D_l)]$, where there are L lakes in the sample. This function is maximized over the set of parameters to be estimated from the data, $\{\delta,\theta,\sigma_1,\sigma_2,\rho\}$, though as discussed below, in our application this parameter set is slightly altered in an attempt to correct for possible endogeneity bias specific to our data.

4. Application of the model

The econometric model is applied to lakeshore property in Vilas County, Wisconsin, a popular vacation destination that has more seasonal than permanent residences. The county has one of the highest concentrations of freshwater lakes in the world, and land development in the county is heavily focused on lake shorelines (Schnaiberg et al. 2002). This study area was chosen because prior work in the region has documented the amenity effects of open-space conservation policies. An analysis of county-level migration rates across the northern forest region – including our study area – found that in-migration rates are higher in counties with more public conservation land (Lewis et al. 2002). A hedonic analysis found that per-foot shoreline property values are higher on lakes with more public land and for which the future development prevented by stricter zoning is relatively high (Spalatro and Provencher 2001). This provides the backdrop to examine whether local public goods that raise the price of land also increase land development.

4.1 Data sources and database construction

In our econometric analysis the unit of observation is the parcel, and the sample consists of only legally subdividable lakeshore parcels in Vilas County. The data were derived from a number of sources, including the GIS database described below, the Wisconsin Department of Natural Resources (WI DNR), USDA soil surveys, and town governments in Vilas County.

Estimating land conversion models requires spatial data for multiple points in time, yet 2003 is the only year for which digitized tax parcel information is available for Vilas County. We therefore developed a method to digitize historical plat maps[6] for Vilas County that backcasts from the 2003 GIS dataset. The resulting dataset enables the consistent tracking of development for all parcels between 1974 and 2003. Hard copy versions of historical plat maps are available for Vilas County in four year intervals from 1974-1998. The method for digitizing the plat maps involves four steps. First, a digital image of the plat maps is obtained from a high resolution scanner.[7] Second, geographic coordinates are assigned to the maps by using a process known as rectifying, whereby coordinates from the 2003 GIS dataset are assigned to control points on the newly-generated digital image of the plat map. After a number of control points are set, the map is assigned the coordinates of the 2003 GIS dataset. In this way, the scanned plat map is now an image file with a distinct spatial location.

The third step is to assign attributes to the parcels of the newly-rectified digital plat maps by working backwards from the 2003 GIS dataset. This process begins by overlaying the 2003 GIS layer with the rectified digital plat map such that parcel boundaries in the rectified plat map are matched to their counterparts in the 2003 GIS layer. Subdivisions are identified where the parcel lines on the 2003 GIS layer do not coincide with any parcel boundary on the rectified plat maps. The last step is to generate a modified copy of the 2003 GIS layer so that it matches the rectified plat map, in effect creating a historical GIS layer corresponding to the year of the plat map. When the lines that delineate a parcel appear in the GIS file but not the digitized plat map of a particular year, the multiple small parcels in the 2003 GIS layer are merged together to represent the *pre-subdivision* parcel. This process is repeated for each historical year that plat maps are available. In the end, each time period—1974 through 1998 in 4 year intervals—has a GIS file with all of the spatial attributes of the parcels.

The database of shoreline development consists of all subdividable parcels on 140 lakes in Vilas County.[8] In 1974, there were 1,310 parcels that could be legally subdivided, and 335 individual subdivisions occured between 1974 and 1998. Approximately 11% of parcels were subdivided more than once. If, after an observed subdivision, a newly created parcel was itself legally subdividable, it was added to the database. Consequently the dataset used in estimation has 1,575 subdividable parcels of land, some of which were not in existence at the start of the study period.

The lakeshore development process in Vilas County is dominated by relatively small developments by many individual landowners, as indicated by the fact that during the study period 82% of recorded subdivisions generated less than six new parcels each. Parcels of more than 1,500 feet of frontage account for only 25% of the recorded subdivisions in the dataset, but generated approximately 49% of all new parcels.

4.2. Open space variables

To examine the role of open-space conservation on shoreline development, we include the following three variables in the analysis: the proportion of a lake shoreline in public conservation land (*Public)*, the average frontage of private parcels on the lake (*AvFront*), which is a measure of current development sparcity (the inverse of development density), and the state of zoning on the lake (*Zoning*). Among these the role of *Zoning* in the analysis requires some explanation.

Throughout the discussion we have alluded to the potential for zoning to preserve open space via limits it establishes on the density of development. For instance, zoning often requires minimum lot sizes. In our application this limit is manifest in a minimum frontage requirement for shoreline properties. During the study period the default minimum was the statewide minimum of 100 feet. Towns were free to exceed the state's requirements, and by the end of the study period 7 of the 14 towns in Vilas County set the minimum shoreline frontage requirement at 200 feet.[9] *Zoning* is a binary variable that takes a value of 0 if the applicable minimum frontage requirement is 100 feet, and a value of 1 if the requirement is 200 feet.

The effect of zoning on the subdivision decision can be divided into a direct effect and an indirect effect. Concerning the former, zoning directly constrains the number of parcels that can be created upon subdivision. This direct effect depends on the amount of parcel frontage, and so in our econometric model *Zoning* is interacted with a parcel's amount of frontage (*Front*). The indirect effect of zoning arises via its effect on the amount of open space preserved on a lake. This effect depends on the level of development at the time the subdivision decision is made; stricter zoning on a lake that is already fully developed under the state minimum frontage requirement of 100 feet will have a much lower conservation effect than on a lake that is relatively undeveloped. Because it impacts the supply of a local public good, this conservation effect may affect *both* the binary decision to subdivide *and* the decision about the number of parcels to create upon subdivision. We capture this conservation effect via the interaction

$AvFront \cdot Zoning$. The greater the value of *AvFront*, the greater the effect of *Zoning* on the conservation of open space.

As discussed earlier, the basic logic that open space may *reduce* the probability of subdivision relies on the complementarity of open space and parcel size. To provide the flexibility necessary to test for such complementarity, in our econometric model we include interactions between our open space variables *Public, AvFront,* and $AvFront \cdot Zoning$, and the parcel's frontage (*Frontage*), which is the relevant measure of a parcel's size.

4.3. Potential endogeneity of open space variables

An important issue in a study of household responses to open space variables is the potential endogeneity of the variables. In our analysis, *Public* and *Zoning* are reasonably modeled as exogenous variables. Almost all public conservation land in the county was acquired in the early 20th century.[10] Widespread logging in the late 1800s – commonly referred to as "the cutover" – cleared most of the original forestland in northern Wisconsin and set the stage for mostly failed attempts to farm newly harvested –and agriculturally marginal –land. Most of the present day tracts of public land in Vilas County were either purchased or forfeited to public control in response to widespread land abandonment in the 1930's-1950's (Flader 1983).

The case for the exogeneity of *Zoning* is equally compelling, because shoreline zoning takes place at the town level, and there are scores of lakes in each town. In addition, contrary to a common lament that zoning ordinances are often ineffective because variances are easy to get – a lament indicating that zoning is to some degree endogenous –lakeshore zoning is apparently enforced in Vilas County: we found that only 5% of the recorded shoreline subdivisions clearly violated the zoning standards at the time of subdivision.[11]

On the other hand, *AvFront* is best modeled as endogenous because it is a function of past subdivision decisions; the same unobservable lake characteristics that led to the current average size of private parcels on the lake may also affect a parcel owner's current subdivision decision.[12] Formally, the endogeneity of *AvFront* arises because our inclusion in the model of the lake-specific random effect (μ_l) introduces a time-invariant spatially-correlated unobservable.[13] The specific econometric challenge is that discrete-choice random effects estimation generates inconsistent estimators if the random effect is correlated with a regressor (Cameron and Trivedi 2006).

We devise two strategies for handling the potential endogeneity of the average frontage of private parcels. First, following prior work on correlated random effects models (e.g. Mundlak 1978; Zabel 1992), we build correlation into the model by specifying the lake-specific effect as a function of each lake's initial average frontage in 1974: $\mu_l = \lambda AvFront_{l,74} + \xi_l$, where $\xi_l \sim N(0, \sigma_1^{\,2})$.[14] The intuition for this specification is that the initial state of development on each lake in 1974 proxies for the unobserved attractiveness of each lake for development. Results from this specification are presented in the next section as model 1; the set of parameters to be estimated is amended to include λ.

The second strategy takes the perspective that we are less interested in the effect on the subdivision decision of *AvFront* per se, but rather in the way the current amount of open space on the lake, as indexed by *AvFront*, modifies the effect of the policy variable *Zoning*. From this perspective a way to deal with the endogeneity of *AvFront* is to simply drop *AvFront* from the model and to specify the effect of zoning on the subdivision decision as a random effect taking the form $Zoning_{lt}(\overline{\beta}^1 + \sigma_{\beta^1}\varpi_l^1) + Zoning_{lt} \cdot Front_n(\overline{\beta}^2 + \sigma_{\beta^2}\varpi_l^2)$ in the Probit model, and

$Zoning_{lt}(\bar{\beta}^3 + \sigma_{\beta^3}\varpi_l^3) + Zoning_{lt} \cdot Front_n(\bar{\beta}^4 + \sigma_{\beta^4}\varpi_l^4)$ in the Poisson model, where ϖ_l^1 and ϖ_l^2 are lake-specific standard normal random variables correlated with one another and, importantly, with the lake-specific effect μ_l, and ϖ_l^3 and ϖ_l^4 are independent standard normal random variables.[15] This random parameters framework accounts for the fact that the effect of zoning on the subdivision decision varies from lake to lake and may depend on unobservable features of the lake, but as opposed to model 1, where the effect is explicitly tied to a lake's development level (as measured by the average frontage of private parcels), it treats the effect of zoning as a random variable possibly correlated with unobserved features of the lake. This random parameters specification is presented in the following section as model 2.

4.4. Other variables used in estimation

Table 1 defines the variables used in estimation and provides summary statistics. Following the conceptual model in section 3, the same variables, including the same interaction terms, are used to predict both the probability of subdivision and the number of new parcels created upon subdivision.

A parcel's frontage (*Front*) is included in the model both in interaction terms for reasons described above, and in a quadratic form. We include two variables depicting soil-related impediments to development: the percent of the parcel that has soil limitations for development (*PSL*), and the percent of the parcel with soil not rated for development (*PSNR*) –an indication of the presence of a wetland. We include the dummy variable *SPLIT* in the belief that although the location of an existing residential structure may negatively affect the potential to subdivide any property, this is especially true of smaller parcels that can be legally split only into two parcels.

Previous hedonic analyses of lakefront property (Spalatro and Provencher 2001; Boyle et al. 1999) have shown that a number of lake characteristics influence shoreline property values.

To account for the effect of such lake characteristics on the subdivision decision we include in the analysis variables concerning water clarity, lake size, lake depth, and distance to the nearest town with major services. Finally, to account for economy-wide fluctuations (e.g. changes in mortgage rates) we include dummy variables for each four-year time interval.

5. Estimation Results

The joint Probit-Poisson model in (17) is estimated with original Matlab code by maximum simulated likelihood using independent sets of 200 Halton draws. Halton draws are a systematic method of drawing from distributions that is useful for reducing the number of simulations while increasing the accuracy of the estimation (Train 2003). The estimation results are generally robust across model 1 and model 2. In the discussion below we focus on model 1.

In model 1 the random effects coefficient for *AvFront* in 1974 (λ) is not significantly different from zero at any reasonable confidence level. Since estimating model 1 without *AvFront* in 1974 produces virtually identical estimates for all coefficients, it appears that the potential correlation between average parcel frontage and unobservable lake-specific effects is not strong. This result is reinforced by the small estimated standard error of the lake-specific effect relative to the estimated constant term in both models. Further, the similarities between estimating the model both with (model 1) and without (model 2) *AvFront* suggest that the econometric problems associated with this variable do not appear to be severe.

The results in table 2 generally conform to expectations and yield the following basic conclusions. First, the probability of subdivision is increasing at a decreasing rate with frontage size. Second, the probability of subdivision is lower for parcels with greater soil restrictions. Third, the expected number of new parcels created upon subdivision is increasing at a decreasing rate with frontage size. Fourth, while the parameter estimates for most parcel-specific characteristics are significantly different from zero, most of the lake-specific characteristics do not significantly impact the probability of subdivision or the expected number of new parcels created upon subdivision.[16] Fifth, there is evidence of unobserved lake and parcel heterogeneity that influences shoreline development decisions, as indicated by the statistically significant values of σ_1 (the standard deviation of the lake effect μ_l) and σ_2 (the standard deviation of the parcel effect ω_n). And finally, unobservables are correlated across the decision to subdivide (the probit model) and the decision about the number of parcels to create (the Poisson model), as indicated by the statistical significance of ρ.[17]

Each of the two open space policies included in the econometric analysis (public conservation land, minimum shoreline frontage requirement) provide the foundation for a pair of tests of whether parcel size and open space are separable, substitutes, or complements in the land value function. The first test concerns the effect of open space on the decision to subdivide, and the second test concerns the decision about the number of new lots created upon subdivision. If parcel frontage and open space are either separable or substitutes in the land value function, then an increase in conserved open space necessarily increases both the probability of subdivision (first test) and the number of parcels created upon subdivision (second test), due to the price premium generated by the conservation of open space.[18] On the other hand, if parcel frontage is a complement to open space, then an increase in protected open space *may* reduce or have no effect on both the probability of subdivision and the number of parcels created upon subdivision.

In our econometric analysis this logic applies in a straightforward manner in the case of open space preserved by public conservation land. We calculated, for each parcel in the sample, the discrete-change effect of a 10% increase in public conservation land. The discrete-change effect accounts for all variable interactions and is the difference in the estimated probabilities with and without the change in public conservation land.[19] The statistical significance of discrete-change effects is calculated by implementation of the Delta Method (Greene 2000). As shown in Figure 1a, we find that the increase in public conservation land has *no effect* on the probability that small parcels are subdivided, and *reduces* the probability that large parcels are subdivided (first test). We also find that the increase in public conservation land has *no effect* on the number of parcels created upon subdivision (second test; results not shown in Figure 1). These results suggest that parcel size and the open space provided by public conservation land are complements in the land value function.

Figures 1b and 1c evaluate the sample discrete change effects of an increase in the minimum frontage requirement from 100 feet (the statewide minimum) to 200 feet (the minimum that applies to many observations in the data set). We find that an increase in zoning strictness generally has no statistically significant effect on the probability of subdivision, and a statistically significant *negative* effect on the number of parcels created upon subdivision for most parcels. Yet this discrete-change effect is not a good measure of the open-space effect of minimum frontage zoning, because, as mentioned in the previous section, such zoning has two effects on the subdivision decision: the *indirect effect* of open space conservation, and the *direct effect* of constraining parcel subdivision; a higher minimum frontage requirement necessarily reduces the number of new parcels that can be created, which in turn reduces the value of subdivision and thus the probability that a parcel in our sample is subdivided. It follows that an increase in zoning strictness may generate an overall reduction in both the probability of subdivision and the number of parcels created upon subdivision, even when parcel size and open space are separable or substitutes in the land value function.

To evaluate the (indirect) open-space effect of zoning, we compare the discrete change effect of zoning on lakes that are relatively developed to those that are relatively pristine. In particular, we compared the effect of an increase in the minimum frontage requirement from 100 feet to 200 feet for a parcel with sample average characteristics for all variables except frontage, on a lake with average development density (*AvFront*=400 feet), and a lake with low development density (*AvFront*=1000 feet). We found that for all parcel sizes examined, the effect of the zoning change on the probability of subdivision was no different for average and low density lakes (first test), and that the effect of the zoning change on the number of parcels created upon subdivision also was no different for average and low density lakes (second test).[20] These results suggest that once again parcel size and open space are complements in the land value function.

6. Landscape and Policy Simulations

While both zoning and public conservation ownership can be used to physically constrain a lake's development to the same long-run level, the approach to the long-run level may be quite different under zoning than under public conservation ownership. To examine the time path under the two conservation strategies, we apply the econometric model in a spatially-explicit landscape simulation model. Following Lewis and Plantinga (2007), we interpret the fitted subdivision probabilities denoted by (13) as a set of rules that govern land-use change. For example, if the subdivision probability is 0.1 for a particular parcel, the owner of the parcel will subdivide 10% of the time if the choice situation is repeated. A random number generator is used to repeatedly draw from a uniform distribution defined on the unit interval and compared with the estimated subdivision probability. If the draw is less than or equal to the estimated subdivision probability for that parcel, the parcel is assumed to subdivide; otherwise the parcel is assumed to remain in its current state. If a parcel subdivides in any given simulation run, the number of new parcels is assumed to be the estimated expectation of new parcels as defined by the Poisson model.[21] Conducting this simulation for every parcel on a lake at each point in time over a long horizon gives a *particular* development path for the lake. Importantly, the simulation embeds the dynamic result that subdivisions in one period affect all future subdivision decisions on all parcels on the lake via the change in the variable *AvFront*. Simulating the development path many times and taking the average across the simulations gives the average development path for the lake.

We evaluate the relative effects of zoning compared to public conservation land by using the landscape simulation model to examine the development path of a hypothetical lake with characteristics that match the average lake in our sample. In particular, we examine the development path under the following policies: i) a baseline policy where zoning sets the minimum frontage of a parcel at 100 feet and there is no publicly-owned shoreline, ii) a conservation policy where the minimum frontage requirement is increased to 200 feet and there is no publicly-owned shoreline (“zoning policy”), and iii) a conservation policy where the minimum frontage requirement remains at 100 feet and 50% of the shoreline is public conservation land (“public ownership policy”). The two conservation policies generate physical restrictions to achieve the same long-run effect: one-half of the shoreline development of the baseline policy.

Figure 2 plots four scenarios for the 200-year development path of a lake with 100,000 feet of shoreline, where each time path represents the average of 100 simulated time paths. In three of the scenarios, the 100,000 feet of frontage is initially divided evenly among 20 parcels, generating a parcel size of 5,000 feet for the baseline and zoning policies, and a parcel size of

2,500 feet for the public ownership policy. To evaluate the effect of the disparity in initial parcel size, we added a fourth scenario in which the public ownership policy is evaluated from an initial position of 10 private parcels of 5,000 feet per parcel. As can be seen in Figure 2, the initial parcel size has relatively little effect on the development path under the public ownership policy.

In the baseline scenario, development grows at an increasing rate for the first 32 years before growing at an increasingly decreasing rate for the remainder of the time horizon. The concavity in growth rates after t=30 appears to be driven primarily by the lower frontage size for the remaining un-subdivided parcels. While the physical restrictions of the two conservation policies (public ownership and zoning) imply a long-run development level equivalent to one-half the baseline, the spillovers associated with these policies result in a lower initial growth rate of development than the baseline. In particular, the public ownership policy generates significantly lower growth in early years than either the baseline or zoning policies. This result is driven primarily by the significantly lower probability of subdivision of large parcels on lakes with relatively high proportions of public conservation land. While the initial growth rate on the lake with the zoning policy is similar to the baseline growth rate, the negative effect of zoning on the density of development keeps the growth rate of development below the baseline in all periods.

7. Discussion

In this paper we examine the effects of minimum frontage zoning requirements and public conservation land on the decisions of private landowners to subdivide and develop lakeshore property. Our analysis provides a number of contributions to the land use literature. First, we empirically model the decision to subdivide and develop large residential parcels and show how this decision problem conceptually differs from the decision to develop an agricultural parcel. Second, we model the joint decision of whether to subdivide a parcel, and how many parcels to create upon subdivision, in a full information likelihood framework that explicitly accounts for sample selection. Third, we develop a random effects framework that accounts for both the panel structure of the data and potentially unobserved spatial heterogeneity. Fourth, we provide econometric tests of the relationship between protected open space and parcel size.

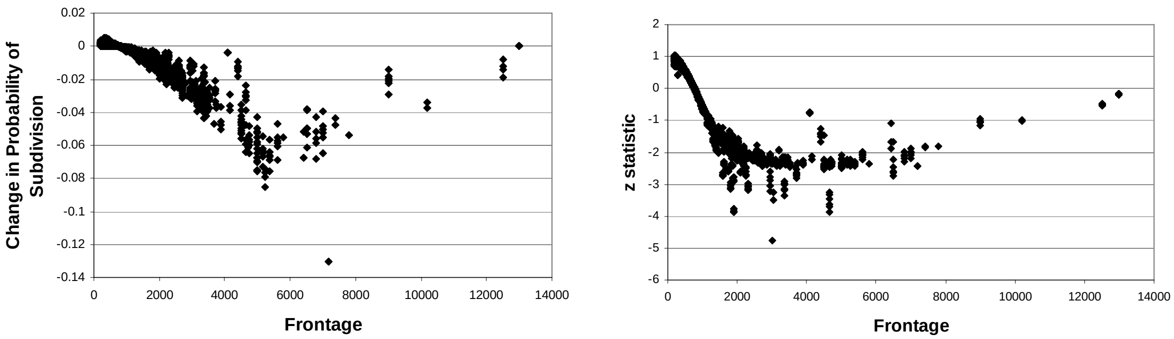
The primary message of this paper concerns the effects of local open space conservation policies on the development of private land. While past analyses have made a compelling case that local open space conservation can increase the development of private land by increasing the returns from development relative to agriculture (e.g. Wu and Plantinga 2003), we find evidence that public conservation land on lake shorelines can actually *reduce* the probability that privately-owned *residential* parcels subdivide and develop. When the development decision is to subdivide and develop a large residential parcel – rather than to develop an agricultural parcel – factors that increase the value of residential land, such as open-space conservation, affect both the returns from developing and the returns from keeping the land in its original state. Theoretically, the effect of open-space conservation policies on the development of existing residential land depends on the relationship –separable, substitutes or complements –between parcel size and local open space in the land value function. Our econometric model provides four tests of the nature of this relationship. The results are consistent with a complementary relationship between parcel size and open space in all four tests, and inconsistent with a separable/substitute relationship in three of the four tests.

Recent results from a survey of lakeshore property owners in our study region shed additional light on this relationship.[22] On average, respondents with large parcels (>500 feet of frontage) reside on lakes with *less* private development and *more* public conservation land than respondents with small parcels. Yet when asked to identify from a list of 18 possibilities the three things that they would most like to change about their lake, 53% of respondents with large parcels, and only 39% of respondents with small parcels, included the change, "reduce the amount of shoreline development". Not only is this result consistent with the conclusion that open space and parcel size are land-value complements, but it invites speculation about why this might be so. One possibility is that open space and parcel size are complements in the landowner utility function. A second possibility is that these attributes are not utility complements, but that the sorting of heterogeneous agents across the landscape effectively generates a complementary relationship in the land value function. This distinction may have significant implications for the design and implementation of open space policy.

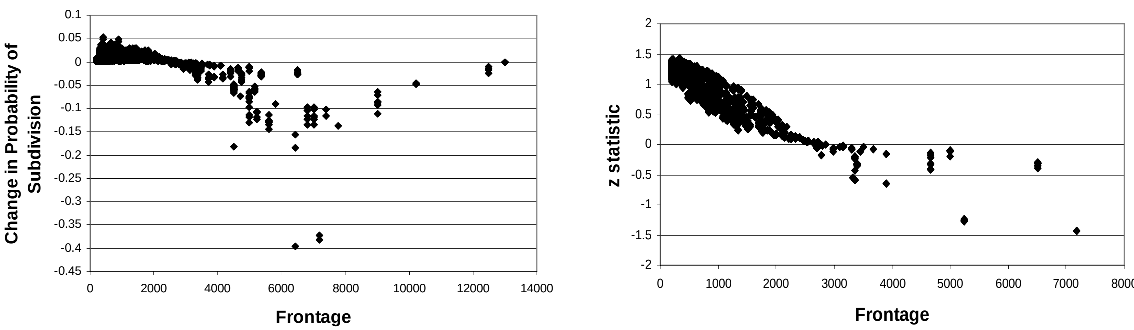
This paper highlights the importance of understanding the spillover effects of local open-space conservation policy on land development. It emphasizes that in many cases the relationship between the creation of local open space and development is complex and not always intuitive. Development density is a key driver of many ecological processes, and so empirical modeling of the effect of open-space spillovers on the spatial dynamics of development can provide important information for conservation practitioners.

Figure 1. Discrete-change effects of open-space conservation policies on probability of subdivision and expected number of parcels across all observations

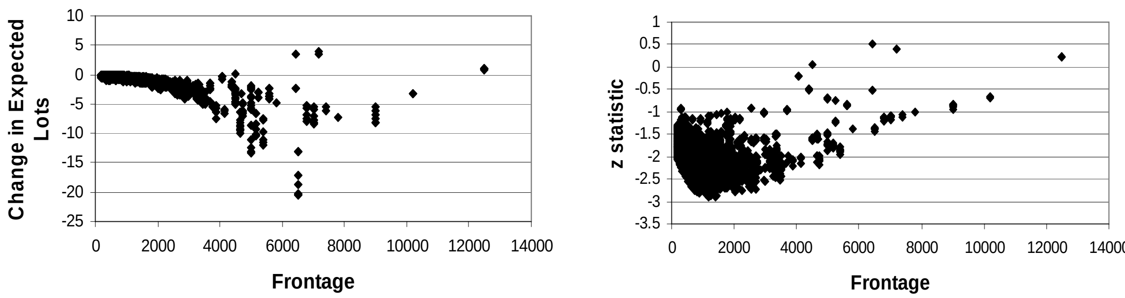
1.a. Discrete-change effects of increasing public conservation land by 10 percentage points on probability of subdivision



1.b Discrete-change effects of 200 ft. vs. 100 ft. minimum frontage zoning on probability of subdivision



1.c Discrete-change effects of 200 ft. vs. 100 ft. minimum frontage zoning on expected number of subdivided parcels



Note: z statistics for all discrete-change effects are calculated with the Delta Method (Greene 2000).

Figure 2. Lake Development Paths with Alternative Open-Space Conservation Policies (average of 100 independent simulated paths on lake with 100,000 feet of frontage)

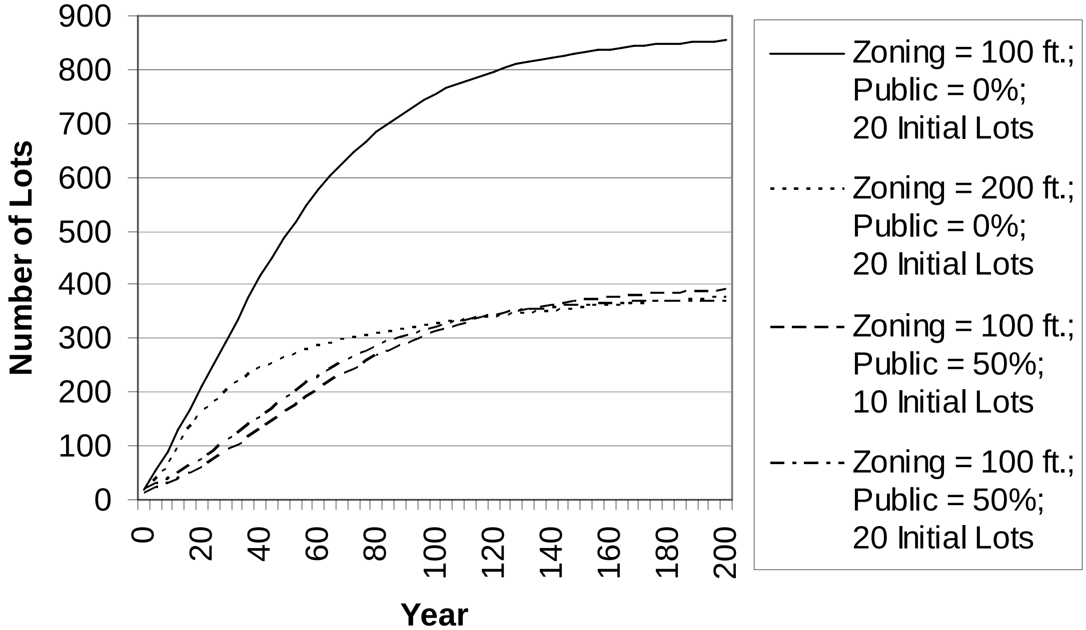


Table 1. Variables Used in Estimation

Variable	Description	Data Source	Average	Min	Max
Parcel-Specific Characteristics					
Front	Shoreline frontage of the property (1000's of feet)	GIS Maps	0.75	0.2	13
PSL	*Percent Soil Limitation:* Percent of the parcel with soil limitations for development	USDA – Soil Surveys	0.56	0	1
PSNR	*Percent Soil Not Rated:* Percent of the parcel with no soil rating (e.g. bog)	USDA – Soil Surveys	0.24	0	0.67
Split	Dummy variable taking a value of one if the only subdivision possible is a split of the parcel;	GIS Maps	0.4	0	1
Lake-Specific Characteristics					
AvFront	Average frontage for all properties on the lake (1000's of feet)	GIS Maps	0.41	0.11	2.4
Public	Proportion of the lake's shoreline owned by County, State, or Federal government	GIS Maps	0.07	0	0.87
Water Clarity	Secchi depth visibility (feet)	WI DNR	6.23	1.23	20.6
Lake Size	Surface area of the lake (acres)	GIS Maps	484	3	3555
Lake Depth	Maximum depth of the lake (feet)	WI DNR	37	3	86
Distance	Distance to the nearest town with major services - Minoqua or Eagle River (km)	GIS Maps	6.47	0.26	171
Zoning	100 ft. (Zone=0) or 200 ft. (Zone=1) minimum frontage zoning	GIS Maps / Townships	0.22	0	1

Table 2. Full-Information Maximum Simulated Likelihood Parameter Estimates

	Model 1				**Model 2**			
	Probit		Poisson		Probit		Poisson	
	Coeff.	St Err.	Coef.	St Err.	Coef.	St Err.	Coef.	St Err.
Constant	-1.50**	0.18	-0.23	0.27	-1.60**	0.17	0.23	0.25
Parcel-Specific Characteristics								
Front	0.42**	0.07	0.92**	0.10	0.45**	0.06	0.68**	0.08
Front ^ 2	-0.49**	0.10	-0.54**	0.12	-0.37**	0.08	-0.46**	0.08
PSL	-0.29**	0.13	-0.29	0.24	-0.35**	0.13	-0.43*	0.22
PSNR	-1.22**	0.21	0.02	0.39	-1.24**	0.21	0.57	0.37
Split	-0.71**	0.09			-0.73**	0.09		
Lake-Specific Characteristics								
AvFront	-0.44	0.27	0.11	0.30				
Public	0.36	0.31	-0.26	0.68	0.39	0.32	-0.82	0.75
Water Clarity	0.01	0.01	0.04**	0.01	0.01	0.01	0.04**	0.01
LakeSize	1.41E-04*	8.E-05	-7.7E-05	1.E-04	1.68E-04*	9.E-05	1.2E-04	1.E-04
LakeDepth	-3.06E-03	3.E-03	-5.5E-03	4.E-03	-3.2E-03	3.E-03	-0.01**	0.004
Distance	0.14	0.09	0.05	0.13	0.11	0.10	0.23	0.12
Zoning					0.32**	0.15	-0.86**	0.21
Interactions								
*AvFront * Zoning*	0.30	0.19	-0.37	0.28				
*Front * AvFront*	0.15**	0.08	-0.15	0.11				
*Front * AvFront * Zoning*	-0.14	0.10	0.16	0.11				
*Front * Public*	-0.51**	0.21	0.09	0.40	-0.58**	0.25	0.57	0.52
*Front * Zoning*	0.02	0.09	-0.17*	0.10	-0.20*	0.12	0.06	0.10
Time-Specific Dummies								
1974	0.03	0.10	0.46**	0.22	0.002	0.10	0.21	0.15
1978	0.09	0.10	0.37**	0.18	0.07	0.10	0.05	0.16
1982	-0.14	0.11	0.40**	0.21	-0.16	0.11	0.20	0.17
1986	0.23**	0.10	0.24	0.19	0.23**	0.10	0.08	0.15
1990	0.05	0.10	-0.01	0.19	0.04	0.10	-0.13	0.18
Random Effects								
σ_1; st. dev. of lake effect (μ_l)	0.22**	0.05			0.24**	0.06		
σ_2; st.dev. of parcel effect (ϖ_n)	0.53**	0.04			0.47**	0.04		
ρ; corr. coef. for υ_{nt}, η_n	0.08*	0.06			0.11*	0.06		
λ; effect of 1974 *AvFront* on μ_l	0.02	0.17						
σ_{β_1}, σ_{β_3}; st. dev. of *Zoning*					0.55	0.45	0.05	0.15
σ_{β_2}; st. dev. of *Zoning·Front*					1.69	2.29	0.42**	0.08
Log Likelihood	-1871.26				-1845.24			

**significantly different from zero at 5% level; *significantly different from zero at 10% level
Note: standard errors for structural probit coefficients calculated with the Delta Method.
Note: off-diagonal Choleski factors in model 2 are not presented.

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Endnotes

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[1] Similarly, there is a set of studies that find lower probabilities of parcel-level development in areas with higher amounts of adjacent development or higher population densities (Irwin and Bockstael 2002; Carrion-Flores and Irwin 2004; Newburn and Berck 2006). In addition, other studies find that higher urban rents increase the probability of land converting from agriculture or forestland to residential development (Lewis and Plantinga 2007).

[2] An exception is Newburn and Berck (2006), who model development as the choice of four density classes in a multinomial random parameters logit framework.

[3] Our dataset may not be that unusual; McConnell et al. (2006) found that only 8% of all subdivisions in an urban-rural fringe region of Maryland developed at the maximum density allowed by zoning.

[4] Typically these models assume a fixed cost of land conversion which is not germaine to our discussion here. Irwin and Bockstael (2004) discuss the conditions under which the sort of simple stopping problem used in the literature is strictly applicable. Although these conditions are strong, the basic model remains generally intuitive and compelling.

[5] Including an inverse Mills ratio as an explanatory variable –the usual method to account for selection bias in ordinary least squares –is not appropriate in Poisson models (Greene 2006).

[6] Plat maps are provided by Rockford Map Publishers, Inc.

[7] When the plat map is scanned, it is simply a picture with no geographic coordinates.

[8] Lakes that were not included either lacked digitized parcel maps in 2003, or a single individual owned the lake.

[9] In our sample, 38 lakes lie in towns where a minimum frontage requirement of 200 feet was in place in 1974, 13 lakes lie in towns that switched to a 200-foot minimum sometime between

1974 and 1998, and the remaining 89 lakes lie in towns that default to the state minimum of 100 feet.

[10] The Northern Highland - American Legion State Forest and the Chequamegon - Nicolet National Forest account for most of the public shoreline in Vilas County.

[11] The particular zoning standard we used to determine violations was the minimum frontage requirement, which is the measure of zoning strictness we use in our econometric analysis.

[12] The endogeneity of neighboring development is discussed as an identification problem by Irwin and Bockstael (2002), and also arises in the more general literature on social interactions (e.g. Manski 2000; Brock and Durlauf 2001).

[13] Ignoring the presence of spatially-correlated errors in discrete choice models will result in inconsistent and inefficient parameter estimates (Anselin 2002).

[14] An alternative approach suggested by Zabel (1992) is to specify the random effect as a function of the average of time-varying covariates.

[15] Correlation between random parameters is achieved by estimating covariance parameters of the estimated distributions as Choleski factors (Train 2003).

[16] Exceptions are that the expected number of lots is greater on lakes with greater water clarity (5% level), and the probability of subdivision is higher on large lakes (10% level).

[17] Significance is determined with a one-tailed test.

[18] This, of course, presumes that open space conservation generates a price premium. Prior hedonic results in our study region find significant residential price premiums on lakes with more public conservation land and stricter minimum frontage zoning requirements (Spalatro and Provencher 2001).

[19] For example, Ai and Norton (2003) found that none of the 72 articles published between 1980 and 1999 in economics journals listed in JSTOR interpreted interaction terms correctly for non-linear models.

[20] Formally, we calculate the following difference: *DC(AvFront=1) – DC(AvFront=0.4)*, where *DC(AvFront=1)* is the discrete change effect of strict zoning on a lake with 1000 feet average frontage and *DC(AvFront=0.4)* is the discrete change effect of strict zoning on a lake with 400 feet average frontage. Standard errors of this difference are calculated with the Delta Method.

[21] The subdivision probability is determined from the version of (13) estimated from model 1, with the lake-specific effect set equal to its mean, $\mu_l = 0$. Conceptually this approach to simulating the subdivision decision is the same as drawing from the distribution of the net benefit function $U(\cdot)$ in (8), and recording a subdivision whenever $U(\cdot) > 0$.

[22] For more details on the survey see www.aae.wisc.edu/provencher .