An Empirical Examination of Real Options and the Timing of Land Conversions

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Introduction[1]

Many studies have examined the effects of land use policies on land prices and urban spatial form. These studies have largely focused on the effects of various land use regulations, including zoning, growth controls, development fees, and differential tax assessment. Theoretical papers have considered their impacts on land prices, development timing, and patterns of land development, and empirical analyses have tended to focus on the effects of these regulatory measures on land prices. While the sizes of the impacts vary, these studies generally find that increased regulation tends to increase house prices and delay development.

Increasingly, jurisdictions have adopted incentive based mechanisms to manage the pace and pattern of urban growth and the conversion of agricultural land. Under one such mechanism, landowners voluntarily receive payment for agreeing to forego conversion and accept easements placed on their land. Since the first 'purchase of development rights' (PDR) program was implemented in 1974, over 53 state and local governments have collectively spent almost $2.4 billion in public funds to preserve nearly 1.5 million acres in the U.S. (American Farmland Trust). In 2002 the Federal government authorized $986 million in matching funds for farmland preservation over the 2002-2006 period. PDR programs enjoy continued taxpayer support; in 2003 alone, $700 million in state and local ballot measures were passed to provide funding for farm and ranch land protection (Trust for Public Land).

In urbanizing areas where landowners can choose to reap immediate financial rewards through development, PDR's offer an alternative that allows them to continue farming while receiving remuneration for their development rights. Empirical studies have characterized decisions to participate in PDR programs (e.g., Nickerson; Duke and others), or evaluated efficiency and distributional aspects of these programs (Nickerson and Barnard; Lynch and Musser), but remarkably few studies have explored the effects of the *existence* of PDR programs on land development decisions themselves. Given the significant costs involved in preserving farmland – which averages approximately $2,000 per acre nationally (American Farmland Trust)

– government agencies are increasingly interested in the effectiveness of PDR programs. Two studies have considered the effects of PDR programs on rates of urban development using aggregate (county level and crop reporting district) data, and find limited evidence that they slow conversions (Miller and Nickerson; Lynch and Carpenter). A few micro-level studies suggest that PDR programs may actually hasten the development of adjacent parcels by making this land more valuable in residential use (e.g., Irwin; Irwin and Bockstael). To our knowledge no studies have explored how the very existence of an option to participate in a PDR program affects landowners' development decisions. That is, even if a landowner ultimately chooses not to preserve, the existence of an option to do so may alter the time at which conversion occurs. Results from real options theory suggest that this may be the case – and, in particular, that the existence of the PDR option may *delay* conversion decisions. If so, these programs may generate benefits (by retaining land in farming longer) beyond those provided by the farmland enrolled in the programs.[2]

Using micro-level data on conversions and preservation of farmland, we empirically test some fundamental results of real options theory in the presence of multiple options. Land conversion decisions lend themselves to real options theory analysis because conversion decisions are irreversible, returns are uncertain, and the decision to convert can be postponed (Titman; Tegene, Wiebe and Kuhn (1999)). Real options theory suggests that the net present value model for investment decisions, which most studies on land use conversion utilize, is incorrect. Uncertainty in returns creates a wedge between the discounted present value of returns and the cost of investment. The size of this wedge is typically increasing in the variance of returns, encouraging postponement of a conversion decision in order to acquire more information. Extensions to the basic theory suggest that faced with multiple options of similar value, a firm or landowner will delay investment decisions (Capozza and Li; Geltner and Riddiough; Trigeorgis).

[1] This work was partially funded by a cooperative agreement with the Economic Research Service of the U.S. Department of Agriculture and by grants from the STAR program of the Environmental Protection Agency.

[2] Both farmland preservation and these additional 'farmland' benefits do come at a cost, however: the foregone benefits associated with development. Whether the one outweighs the other is not at issue in this paper.

Despite theoretical progress on real options, empirical evidence of these effects in the land use context is scant. Quigg (1993) finds that sales prices of land in the Seattle area reflect a premium for the option to delay a development decision. Schatzki (2003) finds evidence that sunk costs and uncertainty in returns lowers the likelihood of land conversion from agriculture to forest in Georgia. Our paper considers how an additional land use alternative, preservation, conveys a different type of option value and how that option affects the optimal conversion time.

Real Options

Three characteristics of investment options explain the observed failure of the net present value rule to characterize investment decisions. First, the option once exercised is irreversible; NPV models implicitly assume that an investment can be reversed if the market is less favorable in subsequent periods, or they assume irreversibility but that the current period is the only period in which the investment can be undertaken. Second, the decision can be delayed. Third, the investment return is uncertain.

The basic real options story is described in Dixit and Pindyck (1994, Ch. 5) and is based on earlier work by McDonald and Siegel (1986). They consider the problem of when to invest in a project with a sunk cost of I and expected return of V, where V evolves according to a geometric Brownian motion:

(1) $$dV = \alpha V dt + \sigma V dz .$$

In (1), α is the 'drift' (i.e. the rate of growth) in expected returns, σ is the standard error of the investment value, and dz is an increment of a Weiner process or the continuous time equivalent of a random walk. Equation (1) implies that the current value of the project is known, but future values are uncertain, are lognormally distributed, and have a variance that grows linearly with the time horizon.[3]

[3] In our model of land conversion, the drift and variance parameters are time varying. This does not change the interpretation.

In our context V is a function of parcel and regional characteristics that are likely to influence development returns. The value of the option to convert land in the future is defined by the function *F(V),* where

(2) $$F(V) = \max_{T} E[(V_T - I)e^{-\rho T}],$$

where T is the time of conversion, I is the cost of conversion including opportunity costs of foregoing future agricultural returns, and ρ is the discount rate.[4] The option will be exercised when the return to investment exceeds the expected capital appreciation.

The solution to the problem must satisfy several conditions, including continuity restrictions and an 'absorbing boundary' condition - if the option value goes to zero it stays zero. Dixit and Pindyck (1994) set of the conditions for solution which defines the optimal return at which investment occurs as,

(3) $$V^* = \frac{\beta_1}{\beta_1 - 1} I,$$

where the term pre-multiplying I represents the wedge between the real options investment rule and the neoclassical investment rule. In terms of the drift and variance concepts from (1), β_1 is defined by:

(4) $$\beta_1 = \tfrac{1}{2} - \alpha/\sigma^2 + \sqrt{\left(\alpha/\sigma^2 - \tfrac{1}{2}\right)^2 + 2\rho/\sigma^2} > 1.$$

This equation yields the comparative static results that are the basis for inclusion of the variance and drift variables in our empirical application. First, an increase in the variance of development returns leads to a decrease in β_1. Since $\partial V^* / \partial \beta_1 = -I/(\beta_1 - 1)^2 < 0,$ this increases the wedge between returns and costs, thereby delaying the decision to develop. Second, as the drift increases, β_1 increases, moving forward the decision to develop.

[4] It is necessary for $\rho > \alpha$. That is, the impatience embodied in the discount rate must exceed the mean increase in return. Otherwise, a landowner would always find it optimal to wait to invest.

Our empirical model investigates a more complex real options problem – one in which land use conversion occurs in the presence of more than one 'investment' option. Specifically, landowners can 'invest' by developing or by selling their rights to develop. Capozza and Li's theoretical model predicts that options to develop at different intensities can delay development. In this paper we test whether a different *land use* option delays development decisions.

Geltner and Riddiough's theoretical work that predicts the decision will be indefinitely delayed when the options are of similar values and the returns are highly correlated. This suggests that landowners might indefinitely delay a development decision if easement values and development returns are both dependent on a single stochastic process – e.g. land prices. In our study area, the easement payment is not necessarily highly correlated with movements in land prices; payment is determined as a static function of parcel characteristics. And, it is the easement payment that is likely to determine whether preservation constitutes a viable option, since a landowner will only preserve if this payment exceeds his reservation price for selling an easement. Thus, the presence of a PDR program in our study area may be expected to delay development decisions on qualifying parcels because of the existence of multiple 'investment' options, but the decision will not be delayed indefinitely.[5]

Hazard model of the timing of land conversion

Many land use studies evaluate conversion decisions utilizing discrete choice models as a function of parcel level attributes (Bockstael; McMillen; Kline and Alig; Landis and Zhang). This approach provides insights on how parcel attributes affect the probability of conversion but does not account for the dynamic environment in which conversion decisions are made. Duration models, on the other hand, are particularly useful for studying factors affecting the occurrence and timing of decisions and are increasingly applied in a land use context (Irwin; Irwin and Bockstael; Nickerson and Bockstael; Hite). We employ duration models because we

[5] A second strand of literature on multiple options deals with how option pricing depends on the nature of the investment, e.g., sequential or conditional investments (Trigeorgis 1998). This strand of literature is less relevant for our current study as we are not trying to price the easement option (for articles on valuing options see, for example, Tegene, Wiebe and Kuhn; Quigg; Plantinga, Stavins, and Lubowski).

are primarily interested in testing predictions of real options theory that affect the timing of landowner decisions to develop.

Duration models are explicitly concerned with the timing of transition from one state to the next, where one mechanism governs both the occurrence and timing of the transition. These models contain two key features useful in our modeling effort: they incorporate the conditional probability that the conversion occurs in period *t*, given that it has not occurred up until period *t,* and they allow incorporation of variables that vary over time – which capture the cumulative effects of changes on the conversion probability.

The distribution of the durations in the undeveloped state can be described in terms of the hazard function or the survival function. First, define *T* as the duration of time the parcel is in the undeveloped state, where *F(t)* is the cumulative distribution of the random variable *T*. The survival function, *S(t)*, is the probability that the parcels survives in the undeveloped state at least to period t, and therefore is equal to *1-F(t)*= Pr(*T*>*t*). The hazard function, *h(t)*, is the probability that the failure event occurs in the time period between *t* and Δt, conditional on the fact that $T \geq t$, i.e., the event had not yet occurred. This function, the hazard rate, is interpreted as the rate at which events occur - in our case, the conversion rate of parcels - and is given by,

(5) $$h(t) = \lim_{\Delta t \to 0} \frac{\Pr\left(t \leq T < t + \Delta t \mid T \geq t\right)}{\Delta t}.$$

The hazard rate is related to the survival function as $h(t) = -d \ln S(t) / dt$ or as:

$$h(t) = \frac{f(t)}{S(t)}$$ where *f(t)* is the density function associated with *F(t).*

Hazard models can be estimated with fully parametric hazard functions as well as semi-parametric (Cox) hazard functions. The principal advantage of the former is that it allows us to estimate the mean duration for ‘censored’ parcels, i.e. parcels that remain undeveloped at the end of the study period. This allows us to determine whether the average time until conversion differs between parcels that face multiple options (development and preservation) and those that face only the development option. A second advantage is that fully parametric hazard functions explicitly deal with censored observations in the likelihood function, unlike the Cox model which estimates a partial likelihood based on the order of conversion events (Cox). Also, Cox models do not handle ties in the data well. Ties are events that occur in the same time period as defined by the periodicity of the data (e.g. conversions that occur within the same year, where the calendar year is the temporal unit of observation). The nature of our data allows us to observe only 11 event times (one per year) and 383 events (conversions) leading to many ties. On the other hand, the main advantage of the Cox model is that it does not require specification of a distribution of the baseline hazard while still providing estimates of the effects of covariates on the hazard rate. It is also considered to be somewhat more robust than parametric models (Allison). We estimate three fully parametric models, as well as a Cox model as a sensitivity check.

The general form for estimating a proportional parametric hazard model is:

(6a) $$h_i(t) = h_0(t)\exp(\beta_1 x_{i1} + ... + \beta_k x_{ik}),$$

where the subscript i denotes an individual parcel and $h_0(t)$ is the baseline hazard that affects all parcels equally. In essence, $h_0(t)$ picks up the development pressure that is being exerted on the study area as a whole which varies over time but not over space. $x_{i1...k}$ are variables affecting the hazard of conversion and $\beta_{1...k}$ are corresponding coefficients to be estimated. Included in $x_{i1...k}$ are factors affecting net returns to development, such as proximity to employment centers and nearby population densities, parcel attributes affecting conversion costs, measures of development regulations, and land quality characteristics affecting returns to agricultural uses. Also included are the 'real options' variables expected to affect the timing of conversion decisions: the variance and drift of expected development returns, as well as a binary variable indicating whether a parcel has the option to preserve land in a PDR program.

The parametric assumption for the baseline hazard enters the model through $h_0(t)$. Taking the log of both sides of equation 6a we have

(6b) $\log h(t) = \omega(t) + \beta_1 x_{i1} + ... + \beta_k x_{ik}$,

where for purposes of subsequent notation, we will define the following:

(6c) $\theta = \exp(x_i \beta)$.

We consider two primary specifications for the baseline hazard which lead to the Weibull and the Gompertz. Both specifications allow the baseline hazard to be a monotonically increasing or decreasing. The Weibull hazard function is given by $h(t) = \theta\alpha t^{\alpha-1}$ and the Gompertz by $h(t) = \theta e^{\gamma t}$. Both of these models are proportional hazard representations because the ratio of hazard functions for two observations is independent of the baseline hazard.[6]

To compare robustness of the coefficient estimates we also estimate a flexible baseline hazard using the piecewise exponential. The hazard function for this distribution is

(7) $h(t;X) = \theta \sum_{m=1}^{M} h_m \delta_m$ where $\delta_m = 1$ for ($a_{m-1} \leq t < a_m$) and = 0 otherwise.

In (7), the a_m's represent a series of breakpoints and the h_m's represent the baseline hazard rates in each of the m intervals. We set our breakpoints such that t_m is a yearly dummy, following the periodicity of our data. This specification allows the baseline hazard to change each year. The key weakness of this specification is the lack of predictive power beyond the last interval in the data. Because we allow h to vary over in a non-systematic way over time, we have no way of predicting h_m for future years.[7]

Parametric specifications are estimated using full information maximum likelihood. For an observation known to experience an event during the study period, the contribution to the

[6] If $\omega = 1$ the Weibull hazard model reduces to the exponential model, a form in which the baseline hazard is constant over time. The nesting of the exponential within the Weibull this allows a likelihood ratio test for model selection between the two.

[7] Extrapolating the h from the final study year into the future turns out to lead to exaggerated estimates for mean duration times of undeveloped parcels in our study.

likelihood function is the value of the density function at time t conditional on the entry time t_0, $f(t,\theta \mid t_0)$. For a censored observation known to survive to time t, the contribution to the likelihood function is the value of the survival function, $S(t,\theta \mid t_0)$, the probability of surviving beyond time t conditional on entry time. Assuming N observations (j = 1,…,N), with U observations that experience an event during the study period ($U \leq N$), the log likelihood function is written as,

$$\ln L = \sum_{j=1}^{U} \ln\{f(t_j,\theta \mid t_{j0})\} + \sum_{j=U+1}^{N} \ln\{S(t_j,\theta \mid t_{j0})\}, \tag{8}$$

where the specific forms of *f(.)* and *S(.)* are determined by the choice of probability distribution for the hazard.

The parameters of the Cox model are estimated by maximizing the log of the partial likelihood function which does not contain the baseline hazard. Unlike most applications of maximum likelihood, each observation in the data set does not necessarily make a contribution to the likelihood function in the Cox model; the ones that do are those that are observed to experience an event during the study period. Information about the censored observations, however, appears in the denominator of each likelihood contribution. The likelihood function is the product of all U contributions. Assuming that exactly one conversion occurs at each event time, the likelihood function is given by:

$$PL(\beta) = \prod_{i=1}^{U} \left[\frac{\theta_i}{\sum_{l \in M(t_i)} \theta_l} \right], \tag{9}$$

where $M(t_i)$ is the set of parcels still "at risk" of conversion at time t_i.

Two issues that arise in using duration models are censoring and exogeneity of time varying variables. Censoring occurs when the study period ends before the last observation 'fails'. In our case we have many right-censored parcels - those that do not develop during the study period. We assume that, conditional on the covariates, the true duration is independent of the starting point (beginning of the observation period), and the censoring time (end of the observation period). This assumption holds when the starting point and censoring time are the same for all individuals, which is the case in our data (Wooldridge (pg. 696)). Time varying variables will be exogenous when they are defined both within and outside the duration *T*. If defined only within the 'spell', they are endogenous (Kalbfleisch and Prentice).[8] Our time varying variables satisfy this exogeneity test.

Study Area and Description of Data

Howard County, Maryland serves as the study area. Located between Baltimore and Washington, D.C., residents commute to both metropolitan areas. This county is under heavy development pressure –in part because of its proximity to two major employment centers, but also because several neighboring counties 'downzoned' in their designated agricultural areas during the late 1970s (to a realized density of no more than 1 house per 15-25 acres). While Howard County's government is concerned with preservation of the county's agricultural heritage, it has not similarly 'downzoned'. All county land outside the public water and sewer service boundaries is in either the rural conservation zone or the rural residential zone, where development is allowed at a realized density of one house per 4.25 acres. Parcels less than 20 acres can be developed on three acre lots. Other than through zoning and offering a preservation option, the county has relied primarily on adequate public facilities ordinances (relating to schools, road capacity, sewer capacity, etc.) to manage the pace and pattern of development.[9]

Unlike the state program in Maryland, Howard County does not require prior enrollment in an agricultural district in order to be eligible to sell an easement. The usual five-year district enrollment period is considered too significant a transaction cost in Howard County's climate of fast-rising land values. To qualify for the county PDR program, a parcel must be at least 100 acres; parcels at least 25 acres qualify if adjacent to at least 50 acres of preserved farmland. In

[8] An example of an endogenous variable arises in the context of a job tenure study using wage on the job as a time varying covariate. When the tenure is complete the wage on that job is no longer relevant.
[9] The adequate public facility ordinance allows the county to postpone, temporarily, new subdivision construction in any planning zone with insufficient school (and, more recently, road) capacity until new infrastructure can be built.

addition, only parcels not served by public sewer and water are eligible. This translates into a requirement that the parcel be in the western part of the county.

The price a landowner can expect to receive for an easement in the county PDR program is based on a published formula. The County pays a higher price for parcels with better soils, more road frontage, more surrounding agriculture, less erosion or drainage problems, and more actively farmed land in the production of food or fiber. The county subjectively ranks the applications based on which parcels contribute more to the farming industry (for example, farms with fertilizer processing equipment or a feed distribution facilities are ranked higher), are most viable and under moderate development pressure. Parcels whose owners have offered to sell their easements are ranked on the basis of the above considerations and the county extends offers until funds are exhausted. Since the Howard County PDR program has been in operation, it has faced severe budget constraints in several years, including inability to fund any purchases in some years. Nonetheless, a considerable amount of land has been preserved within the program. When Howard County instituted their PDR program in 1980, about 34 percent of its 161,408 acres were in farmland. Between then and the end of our study period in 2001, about 16,000 acres were preserved through the PDR program and about 20,000 acres were converted to residential uses. [10]

The data for this study consist of parcel level data for all undeveloped parcels in Howard County, Maryland as of 1990. Pooling several data sources – primarily property tax assessment data and GIS data from the county and state, including actual parcel boundaries – we were able to reconstruct the landscape as of the end of 1990 and identify at what point in time during the 1991 through 2001 study period parcels were converted to house lots.[11] In the final dataset we have included all undeveloped parcels as of the end of 1990 that were eligible to be subdivided

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[10] Howard County landowners also had the option of preserving in the State of Maryland's PDR program. However, the county PDR program had fewer enrollment requirements and paid a higher price per acre for the easements. The differences in transaction costs combined with the lower payments were so pronounced that landowners chose not to enroll in the State program, before and after the county program was active. Thus, only one viable preservation option existed.

[11] These data sources include the Maryland Taxation and Assessment Data for transaction data and most data on parcel characteristics; GIS parcel boundaries, zoning information and sewer service boundaries were provided by Howard County; and census data for tract and region specific variables. GIS data on roads was available from the State Highway Administration and GIS data on natural landscape features from the Maryland Department of Natural Resources.

into 3 or more housing lots.[12] Not all land that was potentially developable at the beginning of the study period had the option to preserve. Only parcels of at least 100 acres were eligible to enroll, unless the parcel was at least 25 acres *and* adjacent to at least 50 acres of preserved farmland. Parcels had to meet minimum soil quality criteria and also could not be located within public sewer boundaries. The final data set included 1,688 parcels totaling 46,000 acres, of which 255 parcels were eligible for preservation at some time during the study period. These data represent a unique and rich view of the land conversion process across space and time in a rapidly developing county.

Important variables affecting the hazard of development include net returns to developed uses. We follow most empirical studies by measuring distance to urban centers to capture the effect of proximity to major employment centers – in our case, Baltimore, MD (*distba*) and Washington, D.C. (*distdc*). Using Census tract data we calculated the number of lots developed in the previous year divided by total number of developable lots at the beginning of the year in each tract (*devRate*) as a proxy for development pressure in the neighborhood of the parcel. Recent empirical evidence indicates that surrounding land uses affect the value of land in developed uses (e.g., Irwin; Irwin and Bockstael; Geoghegan, et al.). Here, we include several variables calculated in units of the percentage of surrounding land use. The surrounding land uses are:

- Commercial, Exempt, and public buildings (*s_comm*)
- Open space (parks, preserved, protected) (*s_open*)
- Undeveloped farm and forestland (*s_undev*)
- Land that has begun the development process (*s_subdiv*)
- Residential land (low, medium, and high density) (*s_residential*).

The surrounding land use measures are calculated from Maryland Department of Planning land use/cover maps and report percent of land within 800 meters of the true parcel boundary. We chose 800 meters because measures calculated using a radii of less than this have not yielded

[12] We defined conversion in this way to avoid counting the development of family lots as a conversion of farmland to residential use.

robust results using data in this study area (e.g., Irwin). These measures are updated annually as neighboring parcels are converted (to commercial or residential uses) or preserved.

To proxy for construction costs we include the percentage of the parcel that is classified as a steep grade (*steep*), the percentage of the parcel that is not suitable for road construction (*notRoadSuit*), the percentage of the parcel that is not suitable for septic (*notSepticSuit*), a dummy equal to one if sewer service does not exist but is planned in the near future for that parcel (*sewerPlnd*), and the construction cost index for the Baltimore region (*cci*). Regulatory variables include number of lots that the parcel can be divided into given zoning regulations (*numlots*), a dummy variable equal to one if the parcel is in an area unconstrained by an adequate public facilities moratorium (*noApfo*), and a dummy variable equal to one if an open space set aside is required (*reqOpenSpace*). Assuming the exogeneity of these variables is reasonable, based on the facts that the Howard County Comprehensive Plan and zoning codes were passed in 1990 and remained intact through our study period.

To capture the effect of returns in an agricultural use on the hazard of development, we include soil measures that reflect the quality of the soil for agricultural purposes. *class1* is the percentage of the parcel in class 1 soils and includes soils that are prime for row cropping. The remaining soil variables *(class2, class3, and class4*) are also measured in percentage terms, and are associated with declining soil quality for agricultural uses. The effect of these variables is measured relative to the worst soils for agriculture (the excluded soil category). Other measures relating to the agricultural use include parcel size (acres and acres squared) and a dummy variable equal to one if the land is being actively farmed, based on whether the parcel receives an agricultural tax assessment (*farmed)*. Although the soil classifications could proxy for agricultural returns, good agricultural soils can also be favorable soils for development, so the expected direction of the effect on the hazard rate is ambiguous. The effect of the size of the parcel on the hazard rate is also ambiguous, as economies of scale may be evident in both farming and development. For a list of variables and descriptive statistics see Table 1.

We also needed to construct drift and variance variables to test our real options theory. As the role of these variables is to capture dimensions of uncertainty in returns to investment (i.e.

returns to development of land for residential use), we used a separate dataset on sales of unimproved housing lots in Howard County from 1986 to 2001, calculating drift and variance variables for each area of the county (defined by Census tract) and each time period. Sales in which price exceeded two standard deviations from the Census tract mean for the year were omitted in order to eliminate the undue influence of outlier parcels whose special characteristics we were unable to identify. After eliminating non-arms length sales and clearly mistyped entries, 9128 observations remained.

Our drift variable for any given tract and year was calculated as the average rate of growth in deflated lot price for sales within the tract over the previous 6 years, corrected for the two principle sources of lot price variation – distance to Washington DC and size of lot. Specifically, for any year, *t*, and tract, *j*, the following OLS regression was estimated:

$$deflatedSP_i = \alpha_0 + \alpha * lagyear_i + \gamma_1 * \ln(distdc_i) + \gamma_2 * \ln(lotsize_i) + \gamma_3 * \ln(lotsize_i^{\,2}) + \varepsilon_i$$

where the observations for the *j,t* regression included all lots sold in tract j during years t-1, t-2, …, t-6. The variable *deflatedSP* is the inflation adjusted sales price in 2000 dollars, *lagyear* equals *s* if the sale took place in year *t-s*, *lotsize* is lot size in acres, and *distdc* is the distance to the center of Washington DC in miles. $\alpha_0, \alpha, \gamma_1,$ and γ_2 are estimated coefficients and the coefficient α on *lagyear* becomes our measure for the drift parameter. Separate regressions are estimated, and therefore distinct drift values are calculated, for each of the 15 tracts and 11 years (from 1991 to 2001) making 165 regressions in all.

The variance was calculated using the unimproved lot sales data for the previous year only. This is the type of information that would be typically available to the average landowner when forming expectations about development returns. The variance measure is defined as

$$\frac{\sum_{i=1}^{N}(deflatedSP_i - \overline{deflatedSP})^2}{N-1},$$

and is calculated for each of the 15 tracts and 11 years in the dataset.

Since we are not concerned with the *level* of variance or drift, each of these measures is standardized by dividing by the mean lot sales price in the respective tract. The average drift for the entire sample is 2.9% and the average standard deviation is 24%. Of course, few observations in a year can lead to a high variance - but this is also a signal of the limited information on recent sales with which current landowners can develop their expectations.

Using this parcel specific, time varying dataset we model the timing of the decision to convert land from an undeveloped (agricultural) state to development. The irreversible decision is defined as the "intent to develop"; this is the failure event in our dataset. Landowners (farmers) can subdivide the land themselves or sell to developers who subdivide the land. During the study period, owners of 383 parcels recorded subdivisions with the county. In general, the time of the conversion event is defined based on the subdivision recording date (for 259 parcels). These included landowners who subdivided themselves and developers who purchased the land and subdivided immediately. However, because the sale of a parcel to a developer marks the intent to develop (and because subdividing sometimes takes time – especially for large subdivisions), we used the date of the prior sale of the parcel if that sale took place within two years prior to the subdivision recording date (for 124 parcels).[13] In our data, the year of most recent sale and the subdivision recording date coincide about 50% of the time. See Figure 1 for the distribution of previous sale dates relative to recording dates.

Finally, in order to test the implications for development timing of the presence of multiple options, we included a binary variable equal to one for parcels that had the option to preserve in a PDR program (and 0 otherwise). This PDR eligibility variable was calculated by applying the county's published eligibility requirements for its PDR program. Because adjacency to already preserved land allowed small parcels ($25 \leq$ acres < 100) to become eligible for preservation, we took account of the fact that the preservation option status could change during the study period.

[13] We are testing the sensitivity of results to different definitions of the conversion time.

Results

Table 2 presents results of the full model under several alternative specifications ranging from the most to the least restrictive representation of the baseline hazard. Instead of reporting the estimated values of the β's from equation (6a), we report the exponential of the estimated β's. Exp(β_k) has an intuitive interpretation. It is the ratio of the hazard evaluated at x_k+1 over the hazard evaluated at x_k, so it equals (the percentage change in hazard)/100 brought about by a one unit change in x_k. To see that this value is independent of other x's and the level of x_k note that

$$\frac{h(x_k+1)}{h(x_k)} = \frac{\lambda \exp[\beta_{-k}' x_{-k} + \beta_k (x_k+1)]}{\lambda \exp[\beta_{-k}' x_{-k} + \beta_k x_k]} = \exp[\beta_k x_k + \beta_k - \beta_k x_k] = \exp[\beta_k]$$

A coefficient estimate less than one implies that the covariate lowers the hazard of conversion and thus delays the conversion time, while those higher than one increase the hazard, bringing forward conversion. For example, the *easement* coefficient is 0.61 in the Weibull model, suggesting that the easement option lowers the hazard rate by 39% (1-0.61).

The exponential hazard is nested within the Weibull specification allowing a test of whether the hazard of conversion is constant over the study period. Likelihood ratio tests for the exponential baseline hazard were rejected at the 1% level versus the Weibull and Gompertz. However, most results do not appear to be very sensitive to the formulation of the baseline hazard, implying that the parametric assumptions on the baseline hazard are not influencing the coefficient estimates for other covariates. While a few coefficients are inconsistent in direction of influence across models, they become insignificant when the direction of influence changes. In what follows, we focus on the Weibull results but note differences across models in important variables.

Variables representing the development pressure for each parcel generally increase the hazard of development, as expected. Parcels in closer proximity to Washington, D.C. have a significantly higher hazard; employment opportunities in Baltimore have an insignificant impact. Conversion of parcels in a planned sewer expansion area is delayed, perhaps due to landowners speculating on the increased land value (i.e. decreased site development costs) when sewer expansion occurs. The surrounding land use variables are all significant, take the expected sign, and, for the most part, are of similar magnitude across models. Parcels surrounded by undeveloped land have a higher hazard, but the effect becomes insignificant in the least restrictive models. That parcels surrounded by more parkland and privately preserved farmland have a higher hazard is consistent with other studies (e.g., Irwin) and suggests that while protection programs (parkland purchases, as well as PDR programs) might preserve individual parcels, they may induce faster development of neighboring parcels.

The effects of regulatory requirements on conversion rates also generally conform to expectations. Not surprisingly, in the absence of an adequate public facilities moratoria, the hazard of conversion increases (although this coefficient loses significance in the less restrictive models). An increase in the lot capacity of the parcel (as dictated by zoning and environmental considerations) leads to an increase in the hazard rate, suggesting economies to scale in development may exist.

As expected, actively farmed parcels (at least according to a definition of farming based on property tax assessment) have lower hazard rates. The insignificance of the coefficients on the soil *class* variables suggests that the opportunity cost of forgoing agricultural returns (which would decrease the hazard) is offset by reduced construction costs (which increases the hazard). Larger parcels have a significantly higher hazard of development.

The primary results of interest are the estimated coefficients on the options variables. We find robust evidence that the addition of a preservation option significantly delays the development decision for qualified parcels. This result is consistent across models. In each case the coefficient on the easement variable is significant at the 95% confidence level.

What is not evident from these models is the impact of the drift and variance variables. The coefficient on the variance in development returns is insignificant, and the coefficient on the drift variable reveals inconsistent directions of influence in the two models in which it has significance. One limitation in our specification is that landowners are assumed to respond similarly to price uncertainty, regardless of the scale of development possible on a given parcel.

To allow for variation in response to price uncertainty across parcels, we re-estimate the models allowing different coefficients for drift and variance according to the parcel's capacity for developable lots. Specifically, the distribution of parcels according to the number of developable lots is divided into deciles and different coefficients for drift and standard error are allowed by decile. [14] These new results are reported in Table 3. Only the results with respect to the option-related variables appear in this table, as all other coefficient estimates remained quite stable when the model was expanded. In particular, the effect of qualification for an easement does not change with this expanded model, still reducing the hazard rate by approximately 40%.

The coefficients on the drift variables remain, for the most part, insignificant. Whether this accurately reflects the failure of this aspect of options theory to apply in the development case or whether the result is due to measurement error is impossible to determine. Table 3 does report some interesting results with respect to the variance of development returns, however. The standard errors are significant for many deciles and exhibit a clear increasing pattern with lot capacity. For parcels that can accommodate few lots (those in the 1st through 3rd deciles, for example), an increase in the variance decreases the hazard and delays development. But for those that can accommodate relatively many lots, an increase in the variance actually appears to move development up in time. The results from the Cox model, which imposes the least structure on the problem, suggests that increases in the variance in returns significantly reduce the hazard rate through the middle percentiles (through the 60th percentile) as well. The results for small developments are consistent with real options theory, but those for large developments do not appear to be. It is possible that the fear of increasing regulation of large subdivisions in the future may counteract the tendency for uncertainty to delay conversion.

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[14] Appendix 1 provides distributional summary statistics for number of lots and their associated acreage for these percentiles. Note that acreage is only partially correlated with number of lots because of variations in zoning regulations across the county. Minimum lot sizes vary significantly and amount of land excluded from development due to environmental sensitivity constrains varies according to parcel characteristics.

How does the presence of an easement option affect the actual timing of the conversion decision? To answer this we predicted the average time until conversion (mean duration) for the censored parcels in the dataset (i.e. those that had not been converted or preserved by the end of the study period). The predicted mean durations are reported by acreage class, starting at 25 acres (the minimum acreage required to qualify for an easement). Results from the Weibull (rather than the Gompertz) model were selected for this exercise on the basis of the Akaike Information Criteria.[15]

Table 4a reports the mean durations for parcels by easement qualification status, using model results corresponding to Table 2, while Table 4b reports the mean durations using the expanded model results from Table 3. In both cases, the predicted mean durations are significantly longer for parcels with an easement option. These mean durations range from about 13 to 18 years (depending on parcel size) for easement-eligible parcels, as contrasted with mean durations of 10 to 15 years for parcels that do not have the easement option.[16] These results suggest that the mere existence of the preservation option extends the time until conversion by 15-25 percent. Also, predicted durations decline as parcels increase in size, but not linearly.

Summary

This study empirically estimates several predictions of real options theory in a land use context, including whether price uncertainty impacts decisions to convert farmland to developed uses and whether the presence of multiple land use options – specifically, an option to preserve farmland in a PDR program – delays development decisions. In doing so we use a duration modeling approach, which captures the conditional dependence of the conversion decision: it explicitly recognizes that the probability that a given parcel will begin the development conversion process in the next period is conditional on the fact it did not convert in any previous period. We find significant evidence that the option to sell a PDR easement decreases the rate of development and does so by approximately 40%. This estimate was robust across different model

[15] The AIC criteria is used for comparing non-nested models, and compares the likelihood values because the same number of parameters are estimated by both models: $AIC = -2(\log \text{likelihood}) + 2(c + p + 1)$,
where c is the number of covariates and p the number of ancillary parameters.

specifications. These findings are in line with real options theory predictions, which imply that the addition of an option to the choice set will increase the value of waiting to the landowner.

An additional element of real options theory suggests that price uncertainty, measured by the variance and drift in development returns, affects the speed of development. We found little support for the contention that drift in returns matters even after adapting the model by allowing different coefficients for parcels with different development capacity. However at least two data-related reasons may account for the non-result. First, the concept of drift makes sense in terms of the theoretical literature on real options, but is difficult to operationalize using real data. Our measure is at best a proxy for the concept and suffers from measurement error from a number of sources. Second, it may be difficult to separate the effects of drift from other time varying explanatory variables such as interest rates, construction cost indices, etc.

With regard to variance in returns, however, we uncovered some interesting results. Increases in variance appeared to affect parcels differently depending on the parcel's development capacity. An increase in variance was found to significantly delay conversion decisions for parcels with lower lot capacity, but tended to speed up conversion for parcels with the highest lot capacities.

We also investigated how long landowners delay conversion decisions, given the existence of an easement option. The existence of an easement option extends the mean duration (delayed the conversion decision) between 2.5 to 3 years, delaying the approximately 15 year average conversion time by 15-25 percent. Given that the induced delay is only a few years, the amenity benefits generated by retaining land in a farming use (which ultimately develops) are probably limited. However, if the delay in percentage terms is found to be consistent across regions, the gains from such programs could be considerably higher in areas where the general speed of conversion is happening at even a slightly slower rate.

Whether PDR programs in other areas do cause a similar delay will likely depend on how similar important program features are to Howard County's program. One important feature is

[16] These numbers are in line with the county planners' expectations of when the county will be fully developed.

how much the PDR program pays a landowner for an easement, relative to what the easement is worth. Howard County has used a static formula to determining the easement price. This design feature may make Howard County's program a less attractive alternative to development than programs whose payment mechanisms more closely follow stochastic land prices. Theory predicts the closer are the values of multiple options and the more correlated their returns, the greater the delay in choosing between options (Gelter and Riddiough). Transaction costs associated with the preservation option may also matter. For example, Howard County's PDR program has not required landowners to enroll in 'agricultural districts' as a prerequisite to selling an easement. District agreements often entail forgoing both development and preservation options for an initial period of several years, while maintaining active farming enterprises. This district requirement can be costly in rapidly developing areas, increasing the wedge between preservation and development option values (i.e., these costs could reduce the value of the preservation option). They may therefore make it easier for landowners to decide between the two options – potentially leading to quite insignificant delaying effects of PDR programs on conversion rates.[17]

This research provides empirical evidence of a previously untested prediction of real options theory: that additional options increase the value of waiting to make irreversible decisions. However, as we found the delay to average only 3 to 5 years, PDR programs may provide only limited additional open space and amenity benefits beyond what is provided on preserved parcels.

[17] Delaying development is a program goal for many agricultural district programs; however, these programs may not delay development at all, if their only participants are those landowners who plan to preserve their land.

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Table 1. Summary Statistics

Variable	Description	Mean	Std.Dev.	Min	Max
	Options Variables				
Drift	Drift in unimproved lot price	2.996192	7.490031	-23.0759	30.14963
Stderr	Standard error of unimproved lot price	24.84786	13.30525	0.05753	60.00807
Easement	Qualified for easement	0.147667	0.354779	0	1
	Returns to Development				
distba	Distance to Baltimore, MD	7.096908	2.666106	0.034306	18.04557
distdc	Distance to Washington, DC	11.85901	2.300788	0.034306	16.99626
devRate	Development rate in census tract of parcels not developed to full potential	5.193862	6.890726	0	62.09935
s_comm	% surrounding land use in commercial	8.99245	10.83037	0	63.90587
s_open	% surr land use in preserved or protect	22.12232	15.00339	0	87.96947
s_undev	% surrounding land use in undeveloped land without a house	10.93621	8.032735	0	56.70102
s_subdiv	% surrounding land use in vacant subdivided land	4.34235	4.524894	0	44.3803
s_residential	% surrounding land use in residential	24.19855	13.17926	0.030071	72.94289
sewerPlnd	Sewer planned	0.134425	0.341117	0	1
Popden	Population density	0.121946	0.15734	0	0.999747
Intrate	Interest rate	4.568636	0.893655	2.9975	5.82
	Construction Costs				
cci	Construction costs index (Baltimore area)	2908.822	201.5257	2508.06	3099.25
steep	% own parcel with steep slopes	0.1554	0.302294	0	1
forest	% own parcel in forest cover	35.90895	36.0916	0	100
notRoadSuit	% own parcel not road suitable	38.41441	27.59166	0	100
notSepticSuit	% own parcel not septic suitable	17.44426	28.26425	0	100
	Regulations and Zoning				
noApfo	No adequate public facilities moratorium (=1 if none, 0 otherwise)	0.912672	0.282323	0	1
reqOpenSpace	Open space required (=1 if yes, 0 otherwise)	0.760141	0.427009	0	1
numlots	Full development potential in number of lots	20.97211	43.91253	3	859.1024
	Returns to Agricultural Use				
farmed		0.195067	0.396263	0	1
class1	% own land use class 1 soils (prime)	1.293702	5.174995	0	66.97584
class2	% own land use class 2 soils	42.23682	30.40452	0	100.0492
class3	% own land use class 3 soils	27.11175	23.36565	0	100.0015
class4	% own land use class 4 soils	9.018492	16.65641	0	99.98521
acres	Parcel size, in acres	27.53602	56.41598	0.779649	1419.976
acres2	Parcel size, squared	3940.824	51567.19	0.607852	2016331

N=1688

Sources: Maryland Department of Planning; Maryland Department of Assessments and Taxation; Maryland Department of Natural Resources; Maryland State Highway Administration; Howard County Department of Planning; U.S. Census Bureau.

Table 2. Hazard Model Estimates

Variable	Weibull Log(L)=-951.90 Ratio	P>|z|	Gompertz Log(L)=-1021.01 Ratio	P>|z|	PW Exponential Log(L)=-1035.21 Ratio	P>|z|	Cox Exact Log(L)=-1536.67 Ratio	P>|z|
				Options Variables				
Drift	+1.011138	0.075	*0.976758	0	0.993593	0.431	0.992888	0.451
Stderr	1.000278	0.958	1.004172	0.433	0.995295	0.357	0.996356	0.5
Easement	*0.611996	0.038	*0.601389	0.031	*0.577294	0.019	*0.569602	0.017
				Development Pressure				
distBA	0.981605	0.508	0.979639	0.459	0.973441	0.323	0.976394	0.41
distDC	*0.923201	0.002	*0.938889	0.012	*0.937283	0.008	*0.937612	0.017
devRate	*1.043397	0	*1.04578	0	*1.033323	0	*1.037677	0
s_comm	*1.018534	0.012	*1.017793	0.016	*1.017787	0.016	*1.01804	0.014
s_open	*1.021331	0	*1.020681	0	*1.020405	0	*1.021541	0
s_undev	*1.015865	0.052	+1.014775	0.066	1.011074	0.174	1.010858	0.223
s_subdiv	*1.039921	0.001	*1.038764	0.001	*1.033243	0.006	*1.034506	0.012
s_residential	*1.026092	0	*1.026385	0	*1.026266	0	*1.027133	0
sewerPlnd	*0.343196	0	*0.347444	0	*0.307815	0	*0.267255	0
Popden	1.10763	0.769	1.108759	0.771	1.012444	0.973	0.966654	0.929
Intrate	0.927379	0.181	*0.905878	0.048	*0.835562	0.023	0.897022	(1)
				Construction costs				
cci	*0.988907	0	*0.995254	0	0.999364	0.217	1.000446	(1)
steep	*0.994762	0.035	*0.994613	0.031	*0.994143	0.018	*0.993734	0.018
forest	*1.003434	0.025	*1.003263	0.033	*1.003106	0.039	*1.003306	0.04
notRoadSuit	*0.986519	0	*0.986559	0	*0.98663	0	*0.985895	0
notSepticSuit	1.001902	0.505	1.001237	0.657	1.001707	0.536	1.002279	0.42
				Regulatory Variables				
noApfo	*1.991308	0.001	*1.646142	0.023	1.069712	0.753	1.08623	0.729
reqOpenSpace	1.249602	0.199	1.241012	0.214	1.232845	0.222	1.277927	0.175
numlots	*1.005111	0	*1.005022	0	*1.005206	0	*1.007268	0
				Agriculture Returns				
farmed	*0.117773	0	*0.114351	0	*0.121164	0	*0.109683	0
class1	1.002731	0.813	1.00177	0.878	1.003915	0.721	1.003296	0.775
class 2	1.001488	0.509	1.001465	0.508	1.002123	0.339	1.002355	0.3
class 3	0.998974	0.69	0.99925	0.768	0.999575	0.867	0.999506	0.856
class 4	1.003346	0.402	1.003759	0.339	1.004362	0.253	1.004325	0.316
acres	*1.025732	0	*1.026757	0	*1.027527	0	*1.026716	0
acres2	*0.999916	0.001	*0.999913	0.001	*0.99991	0.001	*0.999914	0
	p	3.79 (0.2278)	gamma	0.2844 (0.042)				

* significant at the 5% level
\+ significant at 10% level

N=1688

Note: (1) these variables do not vary over time or observations, preventing estimation of standard errors in the Cox model.

Table 3. Selected Estimates from Expanded Model (Options variables)

Note: grp0 represents parcels in the bottom decile in the distribution of lot yields; grp9 are parcels in the top decile.

	Weibull		**Gompertz**		**PW Exponential**		**Cox**	
	Log(L)=-913.66		Log(L)=-986.30		Log(L)=-1000.57		Log(L)=-1503.16	
Variable	Ratio	P>|z|	Ratio	P>|z|	Ratio	P>|z|	Ratio	P>|z|
grp0_drift	1.023283	0.121	.9962025	0.825	1.013243	0.438	1.013596	0.590
grp0_sterr	.9849205	0.121	.9890488	0.281	*.9804963	0.053	+.9811451	0.057
grp1_drift	1.054959	0.108	1.035443	0.378	1.051528	0.168	1.053649	0.138
grp1_sterr	*.9597926	0.001	*.9607075	0.003	*.9521624	0.000	*.9518493	0.001
grp2_drift	.9528924	0.115	*.9251431	0.006	*.9462182	0.049	+.9448401	0.065
grp2_sterr	*.9733181	0.025	*.9769598	0.055	*.9693769	0.010	*.9696668	0.005
grp3_drift	1.007849	0.734	.9740243	0.312	1.002395	0.922	1.002875	0.933
grp3_sterr	+.9802573	0.088	.9842052	0.183	*.9744947	0.033	*.9748558	0.021
grp4_drift	1.015174	0.668	.9841218	0.685	.9987583	0.976	.9990128	0.977
grp4_sterr	.9821466	0.113	.9850808	0.205	*.9754464	0.037	*.9757212	0.020
grp5_drift	1.028864	0.344	.9960801	0.910	1.015173	0.657	1.016028	0.574
grp5_sterr	.9862668	0.153	.9896249	0.316	*.9799292	0.048	*.9803917	0.044
grp6_drift	.9848963	0.573	+.9513613	0.056	.967463	0.213	.9636158	0.142
grp6_sterr	1.001927	0.807	1.006004	0.461	.9974968	0.757	.9990588	0.905
grp7_drift	1.004634	0.749	*.969213	0.046	.9882314	0.475	.9875761	0.559
grp7_sterr	1.005223	0.473	1.008854	0.236	.9994515	0.941	1.00077	0.918
grp8_drift	1.019437	0.172	.981994	0.191	.9998825	0.994	.9997816	0.991
grp8_sterr	*1.015463	0.031	*1.018793	0.009	1.009388	0.169	1.011247	0.144
grp9_drift	1.013282	0.119	*.9691176	0.002	.9856585	0.192	.9821995	0.303
grp9_sterr	*1.039549	0.000	*1.040353	0.000	*1.032123	0.000	*1.035702	0.000
Easementi	+.6364394	0.089	+.6419434	0.088	+.6274188	0.069	.6114643*	0.047

* significant at 5% level
+ significant at 10% level

Table 4a. Predicted Mean Durations by Parcel Size

	Qualified for Easement		
	No	Yes	Difference
Acres (>100)	9.9 (0.61)	13.09 (0.35)	3.19** (0.71)
Acres (75 to 100)	10.85 (0.51)	13.59 (0.45)	2.74** (0.69)
Acres (50 to 75)	11.99 (0.39)	14.46 (0.36)	2.47** (0.54)
Acres (25 to 50)	15.04 (0.26)	17.68 (.038)	2.64** (0.47)

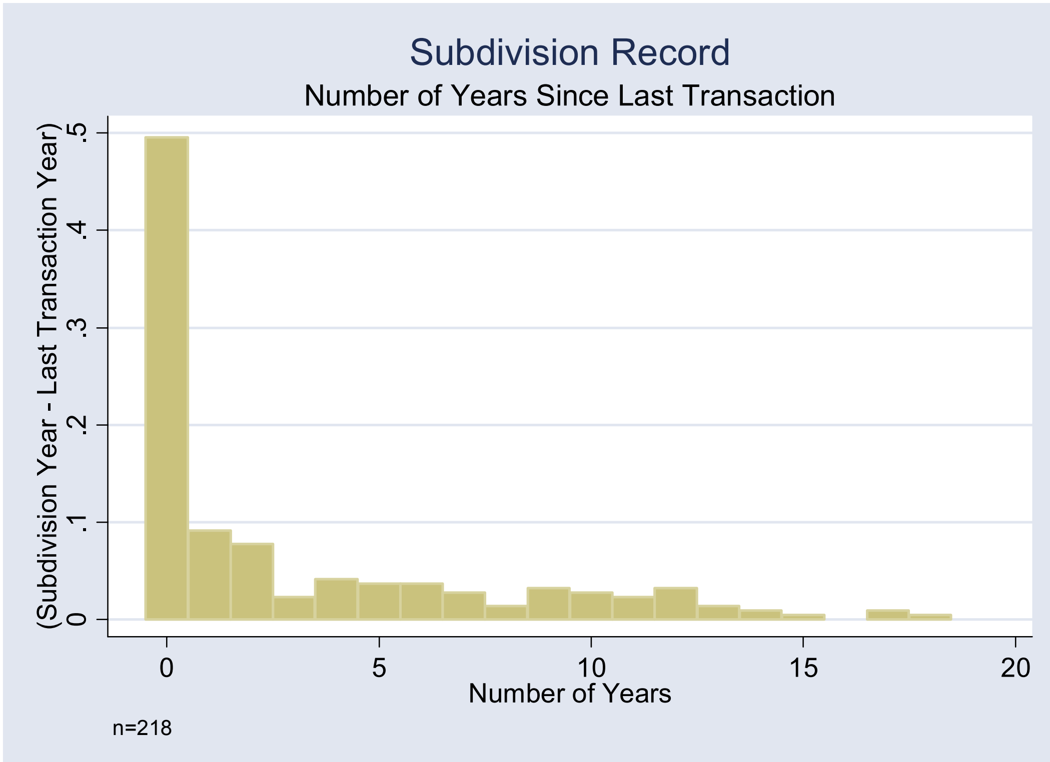
** significant at 1% level. Standard errors in parentheses

Table 4b. Predicted Mean Durations (expanded model) by Parcel Size

	Qualified for Easement		
	No	Yes	Difference
Acres (>100)	10.78 (0.67)	13.18 (0.33)	2.40** (0.75)
Acres (75 to 100)	11.11 (0.52)	13.65 (0.46)	2.54** (0.70)
Acres (50 to 75)	11.88 (0.38)	14.46 (0.35)	2.58** (0.53)
Acres (25 to 50)	15.20 (0.26)	18.28 (0.37)	3.08** (0.46)

** significant at 1% level. Standard errors in parentheses

Figure 1. Distribution of elapsed time between last sales transaction and subdivision recording dates



Source: Interpretation of Maryland Department of Assessments and Taxation.

Appendix 1. Descriptive Statistics, by deciles of distribution of number of lots

Category (percentile)	Variable	Mean	StdDev	Min	Max
0 (lowest decile)	Number Lots	3.298328	0.977534	0.127396	4.506446
	Acres	7.702142	6.626167	0.848223	29.11016
1	Number Lots	5.027978	0.211871	4.511644	5.366228
	Acres	11.8614	10.13895	0.779649	50.35798
2	Number Lots	5.774608	0.231991	5.380042	6.147574
	Acres	13.10127	10.73475	0.958372	49.56123
3	Number Lots	6.521487	0.22899	6.154715	7.017229
	Acres	10.6197	11.21607	1.040862	50.81467
4	Number Lots	7.786827	0.493632	7.01906	8.744633
	Acres	12.14738	15.25849	1.358659	71.92941
5	Number Lots	9.854905	0.712517	8.747434	11.08855
	Acres	17.86175	19.70958	1.46778	69.59932
6	Number Lots	12.69347	1.025953	11.11068	14.51694
	Acres	24.51854	25.63246	1.889206	103.1477
7	Number Lots	17.41767	2.181683	14.5567	22.00215
	Acres	34.50632	35.50078	2.911083	131.9972
8	Number Lots	30.17337	5.863926	22.10401	42.43069
	Acres	67.39617	65.81826	3.771606	198.4414
9 (highest decile)	Number Lots	111.0021	98.23418	42.59061	859.1024
	Acres	75.4709	138.3518	7.267687	1419.976