

AN EMPIRICAL EXAMINATION OF THE TIMING OF LAND CONVERSIONS IN THE PRESENCE OF FARMLAND PRESERVATION PROGRAMS

CHARLES A. TOWE, CYNTHIA J. NICKERSON, AND NANCY BOCKSTAEI

Using a panel of parcel-level data we estimate a hazard model and find strong evidence that the mere existence of an option to preserve farmland delays decisions to convert farmland to developed uses by about six years, a reduction in median conversion time of 12 to 43% depending on parcel size. Where such delays allow local governments to improve infrastructure or implement stricter growth control measures, benefits of a preservation option may be even more long term. Also, increases in the variance of returns to development tended to slow conversion for parcels with all but the highest lot capacities.

Key words: easement, land conversion, proportional hazard model, purchase of development rights, real options, semiparametric duration analysis.

Over the past decades, an increasing number of state and local governments have adopted incentive-based mechanisms in an attempt to manage the pace and pattern of urban growth and the conversion of agricultural land. Under one such mechanism, landowners receive payment for voluntarily 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 fifty-three state and local governments in the United States have collectively spent over \$2.6 billion in public funds to preserve 1.6 million acres (American Farmland Trust 2005a,b). In 2002 the Federal government authorized \$597 million in matching funds for farmland preservation over the 2002–2007 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 2005).

In urbanizing areas where landowners can often choose to reap immediate financial rewards through development, PDR programs offer a means to continue farming while receiv-

ing remuneration for their development rights. Given the significant costs involved in preserving farmland—which averages approximately \$2,000 per acre nationally—government agencies are increasingly interested in the effectiveness of these programs. Two studies have considered their effects on rates of urban development using aggregate (county level and crop reporting district) data and found limited evidence that they slow conversions (Miller and Nickerson 2003; Lynch and Carpenter 2003). A few microlevel studies have suggested that PDR programs may actually hasten the development of adjacent parcels by making this land more valuable in residential use (e.g., Irwin 2002; Irwin and Bockstael 2001). 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. Real options theory suggests 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 even if it is ultimately developed) beyond those provided by the farmland enrolled in the programs.¹

Charles A. Towe, PhD, and Cindy Nickerson, PhD, are economists with the Economic Research Service. Nancy Bockstael is professor in the Department of Agricultural and Resource Economics, University of Maryland, College Park, MD.

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¹ Both farmland preservation and these additional “farmland” benefits do come at a cost, however: the foregone benefits associated with delayed development. Whether the one outweighs the other is not at issue in this article.

In this article we use microlevel data on both the development and preservation of farmland to test whether the option of preserving farmland affects the timing of development. Our model of land conversion decisions is based on real options theory rather than on the traditional net present value rule. We find evidence supporting the theoretical prediction that a PDR program delays development decisions.

Real Options Models

Several authors have recognized that land development is equivalent to the exercise of an option (Dixit and Pindyck 1994; Capozza and Li 2001, 2002). The conditions defining a real option require that the investment is irreversible, that returns are uncertain, and that the decision to convert can be postponed. In contrast to real options theory, the net present value (NPV) rule for characterizing land conversion decisions ignores the implicit costs introduced by uncertainty and irreversibility. It predicts that land will be developed as soon as the present value of development, net of conversion costs, exceeds the present value of the current use. By relying entirely on a one-period rule, the NPV model implicitly assumes that an investment can be reversed if the market is less favorable in subsequent periods.

The real options story recognizes the effects of uncertainty and irreversibility by introducing a value to waiting, as more information emerges. Dixit and Pindyck (1994, Ch. 5) specify a problem in which net return, V , evolves over time according to a geometric Brownian motion as

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

where α is 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. In keeping with the literature, we refer to α as the “drift” parameter and σ as the “variance” parameter (even though σ is actually the square root of the variance). The standard model takes α and σ as constant, but studies suggest that real estate returns are inconsistent with this assumption, at least in the short run (Meese and Wallace 1994; Case and Shiller 1989). Allowing time-varying drift and variance parameters does not change the theoretical predictions, although the value of the option to wait may be lower (Heston 1993).

The NPV rule would predict conversion as soon as $V(t) \geq I(t)$, where $V(t)$ is defined as the value of development in time t minus the lost net revenues due to the nondeveloped use, in perpetuity. $I(t)$ is defined as the infrastructure and regulatory costs of development in time t . Real options theory introduces a wedge, the value of the option to wait, between the net returns and costs. The real options decision rule predicts conversion as soon as

$$(2) \quad V(t) - F(V) \geq I(t)$$

where $F(V)$ is the value of the option. In standard real options theory, the value of the option to wait (i.e., to convert land in the future) is defined by:

$$(3) \quad F(V) = \max_T E[(V_T - I)e^{-\rho T}]$$

where T is the conversion time and ρ is the discount rate.²

Dixit and Pindyck show that the solution to (3), which specifies the optimal development time, is increasing in both α and σ , so that increases in both drift and variance slow development. However, in a real-world setting, the underlying simple assumptions of the Dixit and Pindyck model may not hold. Others have suggested that the fear of preemption may reduce the impact of uncertainty and drive investment decisions back to the standard NPV rule (Williams 1993). Emerging empirical evidence in the real estate market lends support to the notion that greater competition may erode the effect of uncertainty on the timing of development (Shwartz and Torous 2004; Bulan, Mayer and Somerville 2004). Increasing returns to scale may also dampen impacts (Downing and Wallace 2005).

Our empirical investigation deals with a more complex real options problem—one in which land use conversion occurs in the presence of more than one investment alternative. Specifically, landowners can “invest” by developing their parcel or by selling their rights to develop (i.e., selling an easement). The primary goal of the empirical application is to test the hypothesis that the existence of the preservation option delays the development decision.

² A landowner will find it optimal to develop at some point as long as $\rho > \alpha$. That is, the “impatience” embodied in the discount rate must exceed the mean increase in return. Otherwise, a landowner would find it optimal to postpone investment indefinitely.

There is some a priori reason to expect such an effect. Capozza and Li (1994) consider varying time and capital intensity of development options in a real options model context. They show that having variable capital intensity raises the level of the “hurdle” and delays development decisions. Geltner, Riddiough, and Stojanovic’s (1996) work is even more to the point. They model land use choice as a perpetual option where two mutually exclusive types of development (e.g., offices or apartments) are allowed. The authors find that multiple development options delay development decisions and that the more similarly valued the options, the more the development is delayed.

A Hazard Model of the Timing of Land Conversion

Despite theoretical progress on real options, empirical evidence of the aforementioned effects in the land use context is scant. One notable exception is Schatzki (2003) who finds that sunk costs and uncertainty in returns lowers the likelihood of land conversion from agriculture to forest in Georgia. Using a static model, he controls for the presence of multiple options (to convert to urban uses or to pasture) with variables measuring the percent of county land in alternative uses.

Many of the more traditional empirical articles on land conversion decisions, those based on NPV type rules, also use static empirical models. The most common approach is to specify the development decision as a discrete choice (e.g., Bockstael 1996; McMillen 1989; Kline and Alig 1999; Landis and Zhang 1998). This method provides insight into the effect of parcel attributes on the relative probabilities of conversion but does not account for the dynamic environment in which such decisions are made. In contrast, duration models are better able to analyze the timing of the development decision and are increasingly being applied in the land use context (e.g., Nickerson 2000; Irwin 2002; Irwin and Bockstael 2001; Hite, Sohngen, and Templeton 2003).

We use a duration model to analyze whether PDR programs delay development decisions. The duration model can be described in terms of the hazard function. Define T as the “failure” time at which the parcel makes the transition from the undeveloped state to the developed state. The hazard function, $h(t)$, is the probability that the failure event (conversion)

occurs in the time period between t and Δt , conditional on the fact that the failure has not yet occurred by t :

$$(4) \quad h(t) = \lim_{\Delta t \rightarrow 0} \frac{\Pr(t \leq T < t + \Delta t \mid T \geq t)}{\Delta t}.$$

The hazard can be interpreted as the rate at which failures (conversions) occur. Following convention, we specify the empirical model as the natural log of the hazard function:

$$(5) \quad \ln h_i(t) = \omega(t) + x_i\beta$$

where i denotes an individual observation and x_i is a K dimensional vector of variables that are expected to affect the hazard rate and that vary over observations—including the “real options” variables denoting eligibility for preservation and measures of drift and variance. This expression is often parameterized as $h_i(t) = \lambda_0(t) \exp(x_i\beta)$ where $\lambda_0(t)$, which equals $\exp[\omega(t)]$, is the baseline hazard rate. The baseline hazard describes the probability of failure, holding covariates constant, and may or may not vary over time.

Alternative Hazard Function Specifications

When a baseline hazard does not vary over time, it is said to have no duration dependence. Most commonly estimated forms for $\lambda_0(t)$ allow either positive or negative duration dependence, or both, however. In the land conversion case there is reason to expect positive duration dependence (i.e., a hazard that is increasing with time). As time passes, fewer and fewer parcels are available for development because there is a fixed pool of developable parcels in a region and each year some proportion of those parcels drop from the set as they are converted. This suggests the underlying rate of conversion for remaining parcels would *increase* as time passes, as long as demand pressure does not decline. However, there is also an inherent underlying selection process that works in the opposite direction. The parcels with the best unobserved characteristics for development will be the first to be developed, leaving less desirable parcels in the remaining set at risk. Thus, the underlying rate of conversion for remaining parcels may be thought to *decline* because the unobserved quality of the average parcel is declining over time as the best parcels are converted.

We estimate the empirical model three ways. First, we specify the baseline hazard using the

most popular parametric form, the Weibull, which permits a baseline hazard that is a monotonically increasing or decreasing function of time. The baseline hazard is specified as $\lambda_0(t) = \eta t^{\eta-1}$, where η determines the direction of duration dependence. A full parametric specification such as the Weibull uses information on the absolute timing of failures in explaining the rate of conversion over time, allowing the prediction of future “failure.” It also allows the estimation of the mean survival time for “censored” observations, those that have not yet failed by the end of the study period. This is an important feature as a significant portion of our parcels have not converted by the end of the study period.

Second, we specify the baseline hazard as a semiparametric piece-wise exponential. The advantage of this form is its flexibility. It does not impose monotonicity and, in fact, allows the baseline hazard to vary freely from year to year, such that

$$(6) \quad \lambda_0(t) = \sum_{m=1}^M \lambda_m \delta(t)$$

where $\delta(t) = 1$ for $t = m$ and $\delta(t) = 0$ otherwise.

Its chief weakness is its lack of predictive power beyond the last interval in the data: $\lambda_0(t)$ varies in a nonsystematic way over time, so we have no systematic way of predicting it for years beyond the range of the empirical study.

Third, for additional parameter validation, we estimate a semiparametric Cox model in which only a partial likelihood function is estimated, excluding the baseline hazard. The Cox model serves as a sensitivity check, as it is considered somewhat more robust than parametric models (Allison 1995).

Additional Considerations and the Likelihood Functions

In studies with multiple transition options, parcels exiting by an option other than the one of interest require special treatment. In our case, parcels that are preserved during the study period cannot legally be developed, so we treat these parcels as censored. They are removed from the risk set as they are preserved but they are not categorized as “failures” of the type being modeled (i.e., development).

The two parametric specifications (the Weibull and the piece-wise exponential) are estimated using full information maximum likelihood. Contributions to the likelihood

function consist of the probability density function, $f(t, \beta | x)$, for observations that fail during the study period and the survival function, $S(t_c, \beta | x) = 1 - F(t, \beta | x)$, for observations that are censored at time t_c , where $f(\cdot)$ and $S(\cdot)$ are determined by the choice of probability distribution for the hazard. The survival function equals $1 - F(t, \beta | x)$, where $f(t, \beta | x)$ is the cumulative distribution function, and the hazard function is related by the expression $h(t, \beta | x) = f(t, \beta | x) / [1 - F(t, \beta | x)]$.

Thus the likelihood function is given by:

$$(7) \quad \ln L = \sum_{j \in U} \ln \{h(t_j, \beta | x_j)\} + \sum_{j \in C} \ln \{S(t_{cj}, \beta | x_j)\}$$

where U denotes the set of observations that fail during the study period and C denotes the set of observations that are censored. For most censored observations, t_c is the end of the study period, but for those that are removed from the risk set because of preservation, t_c is the year the parcel preserves.

The parameters of the Cox model are estimated by maximizing the log of the “partial likelihood” function, which includes contributions only by those observations that fail during the study period. Information about all observations at risk at each failure time, including the censored observations, appears in the denominator of each likelihood contribution. Assuming that exactly one failure (conversion) occurs at each event time, the partial likelihood function takes the form:

$$(8) \quad PL(\beta) = \prod_{j \in U} \left[\frac{\exp(x_j \beta)}{\sum_{k \in R(T_j)} \exp(x_k \beta)} \right]$$

where $R(T_j)$ is the set of parcels still “at risk” of conversion at T_j , the time the j th parcel converts. In our application the event time is measured in years. As more than one failure occurs in each year, “ties” are handled using the Efron method (Kalbfleisch and Prentice 1980).

The Study Area

To test aspects of real options theory as it applies to land conversion, and in particular to test whether the existence of a preservation option has an effect on the timing of development, we use data on preservation and

development decisions in Howard County, Maryland. The eastern portion of this 160,000-acre county is heavily developed, often at high densities, although many developable parcels still remain. Farmland can be found mainly in the western portion, but this is also the location of many low-density developments. Situated as it is between Washington, DC to the south, Baltimore, MD to the east and Frederick, MD to the west, the entire county is subject to considerable development pressure. This has been especially true since the late 1970s when several neighboring counties “downzoned” their designated agricultural areas to extremely low densities of not more than 1 house per 15–25 acres. In Howard County all developable land outside the public water and sewer service boundaries—and therefore all land nominally eligible for preservation—can be developed at a density of one house per 3 to 4.25 acres, depending on parcel characteristics.

Growth management tools used in Howard County include zoning regulations that specify differential minimum lots sizes and set asides and adequate public facilities moratoria that temporarily close growing areas to new development while infrastructure is built. The county’s purchase of development rights program also helps the county manage growth by giving landowners an alternative to developing farmland. Between its inception in 1980 and 2001, the program enrolled about 16,000 acres—almost 30% of the 1980 farmland inventory and 10% of the county’s overall acreage. Over the same twenty-one years, 20,000 acres of farmland were developed.

Many features of the Howard County PDR program are common to programs in other localities: landowners are paid to accept easements that prohibit conversion of their land to nonagricultural uses; easements are attached to the land with restrictions applying to all current and future landowners; and easements are long term (perpetual in our study area). Details of eligibility criteria, selection mechanisms, and easement pricing vary considerably, however. In Howard County’s program, an agricultural parcel is eligible for preservation if it can legally be subdivided and constitutes at least 50 acres (or at least 20 acres if adjacent to other preserved land or parks). The parcel must also meet minimum soil criteria.

Both selection criteria and easement price are based on a point system that increases with parcel size, percentage of high-quality soils, road frontage, adjacency to other protected land and other agricultural land, suit-

ability for sustained agriculture, history of sustained farming activities, presence of significant natural resources, and absence of erosion or drainage problems. These factors figure into a published formula by which easement price per acre can be calculated for any given parcel. The price per acre varies over parcels depending on characteristics, up to a maximum of \$20,000/acre during our study period. Most of the uncertainty surrounding the preservation option has arisen, not because of easement price, which has been essentially predetermined by the formula, but because of the program’s budgetary limitations, which restricted the number of parcels that could be preserved in any year and effectively precluded any preservation in some of the study years.

The Data and Variables

The data for this study are available at the parcel level and include information about all undeveloped parcels in Howard County 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 reconstructed the spatial landscape as of 1990, as well as the history of parcel conversion from 1991 through 2001. The final data set includes all undeveloped parcels as of 1990 that were eligible to be subdivided into three or more housing lots. We define conversion in this way to avoid counting the development of family lots as a conversion of farmland to residential use. The conversion (failure) time is defined as the date the lots of a subdivision are legally recorded.

Of those parcels that were potentially developable at the beginning of the study period, only those meeting the eligibility criteria had the option to preserve. Because the size requirement for eligibility is reduced when adjoining parcels preserve, 25 parcels that were ineligible in 1990 became eligible during the study period. The final data set includes 1,688 parcels totaling 46,000 acres. Of these 383 parcels (14,545 acres) developed during the study period; 255 parcels were eligible for preservation at some time during the period and 61 of these (6,692 acres) were preserved in the PDR program by 2001.

Equations (1) through (3) contain the theory that informs our empirical model. As with the NPV investment rule, factors that increase returns net of opportunity costs (V) or

decrease investment costs (I) will make a parcel more profitable for development. But unlike the NPV rule, real options theory suggests that an option value exists that drives a wedge between V and I . This “risk premium” is expected to be a function of the drift and variance in net returns and the number of alternative options available to the decision maker.

In the context of housing development, V represents a one-time net return, the price of the developed lots sold to households, minus the present value of the foregone stream of earnings from the undeveloped use. Defined in this way V will be a function of parcel and neighborhood characteristics that are likely to influence the value of the housing lots to consumers, regional factors associated with demand pressures for new housing, as well as physical features and market forces affecting agricultural returns.

Among the most commonly considered factors in V are commuting costs to major employment centers—in our case, Baltimore, MD ($distBA$) and Washington, DC ($distDC$). Recent empirical evidence indicates that surrounding land uses also affect the value of land in developed uses (e.g., Irwin 2002; Irwin and Bockstael 2001). We aggregate surrounding land uses into five categories: residential, commercial/exempt, protected land, roads, and developable land (the normalized category). The surrounding land use measures are calculated as percentages of land within a 100-meter buffer around the true boundary of each parcel.³ These measures are updated in each analysis year as neighboring parcels are converted or preserved.

We include a variable ($popDen$) that measures density of housing at the Census tract level. This serves as a proxy for important surrounding landscape attributes (such as congestion) at a scale larger than the immediate neighborhood. To account for some of the many other factors that vary over regions of the county and over time (such as availability of services) that affect localized demand for housing, we include the rate of development in the Census tract in the previous year ($devRate$).

Zoning regulations affect returns by specifying the maximum number of lots that can

be subdivided ($numLots$) and any open space set-aside requirements ($reqOpenSpace$) that might pertain to the parcel. We treat these regulations as exogenous, given that the Howard County Comprehensive Plan and zoning codes were passed in 1990 and remained unchanged through the study period.

To capture opportunity costs and more specifically the effect of returns in an agricultural use, we include soil measures that reflect the quality of the soil for agricultural purposes ($class$ 1 through $class$ 4). The effect of these variables is measured relative to the worst soils for agriculture (the excluded soil category). Although the soil classifications could proxy for agricultural returns, good agricultural soils can also be favorable for development, making the expected effect on the hazard rate ambiguous. Another measure relating to agricultural use is a quadratic in parcel size ($acres$ and $acres^2$). The likely effect size has on the hazard rate is also ambiguous, as economies of scale may be evident in both farming and development. As an indicator of another type of opportunity cost, the binary variable ($hasHouse$) is included and equals one for parcels that already have an existing house, even though the parcel is zoned such that it can accommodate subdivision into at least three additional lots. We expect lower development probabilities for these parcels as the opportunity cost of development may now include amenity values such as recreation or privacy if the owner resides on the parcel.

In the model of land conversion, I represents the cost of developing the subdivision. To proxy for construction costs we calculate measures of parcel slope ($steep$), forest cover ($forest$), road suitability ($notRoadSuit$), and septic suitability ($notSepticSuit$). We also include a dummy variable equal to one if sewer service does not exist but is planned in the near future for that parcel ($sewerPlnd$). The prime rate ($intRate$), which varies only over time and not parcels, is included as an indicator of the cost of carrying the land from the time the development process is initiated until the lots are sold.

Finally, adequate public facilities moratoria were imposed on development activity in some planning areas in some of the study years. We include a variable ($Apfo$) equal to zero, one-half, or one for any year in which the parcel is in a planning area constrained by an adequate public facilities moratorium relating to school capacity for none, half, or all of the year, respectively. This variable is updated each year of

³ We also tested the sensitivity of results to the use of 400 meter or 800 meter buffers around parcel boundaries. Results on the options variables, the primary variables of interest, were not sensitive to use of these larger radii but some of the surrounding land use measures were.

the analysis as adequate public facilities moratoria are introduced and phased out.

Real Options Variables

Our primary interest is whether the presence of the easement option delays development. The variable (*Easement*) equals one in the years a parcel is eligible to sell an easement and a county budget exists to purchase easements. The variable is updated each year so as to include parcels that become eligible during the study period.

Our secondary interest is in the other predictions of real options theory—whether drift or variance or both has an effect on conversion timing. Empirical testing of the theory is challenging because of the difficulty of measurement; and few precedents in measurement can be found in the literature. In essence we need empirical measures that capture expectations of future trends and variability in net returns for any given parcel at each point in time, and these measures need to vary over parcels for us to have any chance of detecting their effect.

One measurement option is to adopt a rational expectations model, but this is problematic as subdivision activity can have a sufficiently large effect on housing prices to make endogenous any expectations variable based on future prices. We take a more traditional approach and calculate expectations variables as functions of recent market history. Using a separate data set on sales of new housing lots, defined as lots that were built on within ten years prior to the sale, we calculate different drift and variance variables for each of 15 tracts or groups of tracts in the county and for each year of the analysis (eleven years from 1991 through 2001).⁴ Sales in which price exceeded two standard deviations from the Census tract average for the year were omitted in order to eliminate the undue influence of outliers whose special characteristics we were unable to identify. After eliminating these outliers, nonarms length sales, and clearly mistyped entries, 14,998 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 three previous years, corrected for the principal sources of price variation—distance to Washington, DC lot size, square footage of house, and quality of construction, represented by the \mathbf{Z} vector in equation (9) below. For example, we viewed the landowner/developer as forming expectations on the drift and variance in returns from an investment in 1996, on the basis of housing sales in subdivisions within the relevant Census tract over the years 1993, 1994, and 1995.

Two regressions for each year of the analysis are estimated, one for the designated rural area of the county (5 Census tract groups) and a second for the urbanized areas (10 Census tract groups), where each regression includes all qualifying sales for years $t-1$, $t-2$, and $t-3$, where $t = \{1991, \dots, 2001\}$.⁵ Thus the impact of \mathbf{Z} on price is allowed to vary across the rural/urban designation within the county. Specifically, for any analysis year t and Census tract j , the following OLS regression was estimated:

$$(9) \quad \ln(\text{deflated}SP_{j,i}) = \delta_{tj} + \alpha_{tj}s_i + \gamma'_{tj}Z_i + \varepsilon_i$$

where the t, j th regression includes all sales in Census tract j in years $t-1$, $t-2$ and $t-3$. The variable $\text{deflated}SP_i$ is defined as the inflation-adjusted sales price in 2000 dollars for the i th sale; $s_i = m$ if the sale occurred in time $t-m$. The estimated coefficients δ and α vary over tracts and analysis years, but we constrain the γ 's to be constant across Census tracts within each of the two large categories of tracts—rural or urban ($J = 1$ or 2)—in order to conserve degrees of freedom. Given the eleven years of analysis and the 15 Census tracts, 165 different values of the δ 's and α 's are estimated.

The coefficient α_{tj} becomes our measure for the drift parameter for the j th Census tract and the t th year of analysis. The variance measure is a function of the sum of squared residuals. For Census tracts j and analysis year t , it is defined as:

$$(10) \quad \frac{\sum_{i \in L_{tj}} (\text{deflated}SP_i^{\wedge} - \text{deflated}SP_i)^2}{\ell_{tj} - 2}$$

⁴ We refer to the subareas as tracts throughout the article, but in reality some subareas include more than one neighboring tract because of the tract size and limited sales activity. We aggregate some tracts in order to have enough sales in each subarea when estimating drifts and variances, while keeping the size of each subarea as similar as possible to that which is most likely to be considered by landowners when forming expectations.

⁵ Throughout we have ignored the possibility of spatial autocorrelation. If present in a continuous regression model such as our housing price model, the standard errors of estimates will be biased. However, the coefficients and residuals—which form the basis for our estimates of the drift and variance parameters—remain unbiased.

where L_{ij} is the set of all sales in Census tract j in years $t-1$, $t-2$ and $t-3$ and ℓ_{ij} is the number of observations in that set.

The variable $\widehat{\text{deflatedSP}}_{in}$ is the variance-adjusted expected sales price, calculated as $\exp\{\widehat{\text{deflatedSP}}_{in} + 0.5[\epsilon_i^2/(\ell_{ij} - k)]\}$ where k is the number of regressors. Since we are not concerned with the absolute level of variance, this measure is standardized by dividing by the mean sales price in the respective tract. We take the square root, to be consistent with the literature. The calculation is made for each of the 15 tracts and 11 years in the data set.

The average drift for the entire sample is 0.32% and the average variance parameter is 12.9% (with standard deviations of 5.28 and 6.24, respectively). Of course, a small number of observations in a year can lead to a high variance—but this is also a signal of the limited information on recent sales from which current landowners can develop their expectations.⁶

For a list of all variables in the empirical model and their descriptive statistics, see table 1.

Results

Table 2 presents results of the full model under three alternative specifications ranging from the most to the least restrictive representation of the baseline hazard. For each coefficient, we report the *exponential* of the estimated β_k , because $\exp(\beta_k)$ has an intuitive interpretation. It is the ratio of the hazard evaluated at $x_k + 1$ over the hazard evaluated at x_k , where x_k is the explanatory variable associated with coefficient β_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 exponential of the coefficient on the *distDC* variable averages 0.98 over all model specifications, suggesting that the effect of an increase of 1 mile in commuting distance to Washington, DC lowers the hazard rate by 2%.

Likelihood ratio tests for the piece-wise exponential baseline hazard versus the Weibull were rejected at the 1% level, suggesting the hazard of conversion is not constant over the study period. However, most results do not

appear to be very sensitive to the specification, implying that the parametric assumptions on the baseline hazard are not unduly influencing the coefficient estimates for other covariates.

Consistent with other studies, variables representing development pressure, opportunity costs and construction costs generally affect the hazard of development as expected. The coefficients on the surrounding land use measures are all positive and significant relative to the omitted category (developable land). Surrounding protected lands and roads exhibit the largest effects. With an added 1% of surrounding land dedicated to protected areas, the hazard rate of development increases by 3.2–3.6%, suggesting that neighboring *permanent* open space raises the value of housing developments relative to neighboring space that is itself eligible for development. An increase in the percentage of surrounding land dedicated to roads also increases the development hazard, but this effect may be due both to the appeal of better access and to the decreased construction costs associated with more existing road frontage. An increase in the lot capacity of the parcel leads to an increase in the hazard rate, suggesting economies of scale in development. Open space requirements appear to hasten development, consistent with the notion that permanent open space, this time as part of the development, confers positive amenity value and increases the value of developable land (Hardie, Lichtenberg, and Nickerson 2007).

The insignificance of the coefficients on most of the soil class variables suggests that the opportunity cost of forgoing agricultural returns (which would decrease the hazard) is generally offset by reduced construction costs (which increase the hazard). The hazard of development also increases in parcel size (even holding number of lots constant), but does so at a decreasing rate. And, consistent with expectations, a development-eligible parcel with an existing house is at least 78% less likely to be developed than open land.

As expected, location within a planned sewer expansion area lowers the hazard rate by at least 64%, suggesting a tendency to postpone development until cost-reducing sewer expansion occurs. Adequate public facilities moratoria lower the hazard of development, but the effect is not statistically significant. This weak result may be due to our inability to capture the workings of adequate public facilities moratoria in a simple variable.

⁶ Our specifications assume a constant and homogeneous discount rate (ρ) across landowners.

Table 1. Summary Statistics

Variable	Description	Mean	S.D.	Min.	Max.
Returns to Development					
<i>distBA</i>	Distance to Baltimore, in miles	17.969	6.842	6.770	51.325
<i>distDC</i>	Distance to DC, in miles	29.839	5.659	12.947	42.739
<i>sluResidential</i>	% surrounding land use in housing	27.777	21.615	0.000	94.235
<i>sluCommExempt</i>	% surrounding land use in commercial/ institutional	4.756	9.824	0.000	77.233
<i>sluProtected</i>	% surrounding land use in park, preserved, or protected status	14.785	17.095	0.000	92.613
<i>sluRoad</i>	% surrounding land use in roads	8.662	8.831	0.000	57.880
<i>popDen</i>	# households per acre in census tract, 1990	0.528	0.459	0.069	2.363
<i>devRate</i>	Previous year development rate in census tract	5.194	6.891	0	62.099
<i>numLots</i>	# lots allowed per zoning regs	20.999	43.901	3.000	859.102
<i>reqOpenSpace</i>	= 1 if open space required, 0 if no	0.758	0.428	0.000	1.000
Opportunity Costs					
<i>class1</i>	% of parcel with class 1 soils (prime)	1.294	5.175	0.000	66.976
<i>class2</i>	% of parcel with class 2 soils	42.237	30.405	0.000	100.049
<i>class3</i>	% of parcel with class 3 soils	27.112	23.366	0.000	100.002
<i>class4</i>	% of parcel with class 4 soils	9.018	16.656	0.000	99.985
<i>acres</i>	Parcel size, in acres	27.536	56.416	0.780	1419.976
<i>hasHouse</i>	Existing house on parcel	0.992	0.090	0.000	1.000
Construction costs					
<i>steep</i>	% of parcel with steep slopes	15.540	30.229	0.000	100.000
<i>forest</i>	% of parcel in forest cover	35.909	36.092	0.000	100.000
<i>notRoadSuit</i>	% of parcel not road suitable	38.414	27.592	0.000	100.000
<i>notSepticSuit</i>	% of parcel not septic suitable	17.444	28.264	0.000	100.000
<i>sewerPlanned</i>	Sewer planned in next 10 years	0.134	0.341	0.000	1.000
<i>intRate</i>	Annual prime rate	4.569	0.894	2.998	5.820
<i>Apfo</i>	= 1, 1/2, or 0 if restricted by adequate public facilities moratoria for full, 1/2, or no part of year	0.099	0.259	0.000	1.000
Options Variables					
<i>Easement</i>	Qualified for easement	0.108	0.310	0.000	1.000
<i>Drift measure</i>	Drift in sales price	0.323	5.285	-18.508	15.656
<i>Variance measure</i>	Standard error of sales price	12.031	6.239	5.551	31.176

Note: Data sources include Maryland Department of Planning; Maryland Department of Assessments and Taxation; Howard County Department of Planning; U.S. Census Bureau.
N = 1,688 parcels × years = 18,568.

Drift and Variance Real Options Results

Findings with respect to the options variables measuring uncertainty in returns to development, reported in table 2, are not particularly convincing. There is no statistically significant evidence that variance affects development. And the evidence for a drift variable effect is weak, at best. If there is any significant effect, increasing drift appears to hasten development, but only at the 10% confidence level and only for the two more flexible specifications. A positive effect is not consistent with conventional real options theory but is consistent with the notion that competition can induce developers to act in periods of both expected price increases and decreases (Grenadier 2002).

Given the difficulty of translating real options theoretical concepts into empirical measurements, weak results are hardly surprising. A further potential limitation of the specification we have proposed thus far for these real options variables is that landowners are assumed to respond similarly to uncertainty, regardless of the scale of development possible on their parcels. Most real options models are solved for one type of production technology, in particular those that are characterized by decreasing or constant returns to scale (Downing and Wallace 2005). Yet, our finding that more allowable lots, *ceteris paribus*, increases the hazard, suggests increasing returns to scale in development. To allow for variation in response to uncertainty, we divide the distribution of parcels into quintiles

Table 2. Hazard Model Estimates

Variable	Weibull Log(L) = -1001.662		PW Exponential Log(L) = -1012.998		Cox Exact Log(L) = -2553.088	
	Ratio	<i>p</i> > z	Ratio	<i>p</i> > z	Ratio	<i>p</i> > z
Returns to Development						
<i>distBA</i>	0.993	0.465	0.996	0.702	0.999	0.788
<i>distDC</i>	0.980*	0.049	0.981*	0.058	0.981*	0.080
<i>sluResidential</i>	1.001*	0.021	1.011*	0.009	1.012*	0.002
<i>sluCommExempt</i>	1.021*	0.000	1.021*	0.000	1.022*	0.000
<i>sluProtected</i>	1.034*	0.000	1.032*	0.000	1.036*	0.000
<i>sluRoad</i>	1.032*	0.000	1.030*	0.000	1.032*	0.000
<i>popDen</i>	1.050	0.743	1.025	0.865	1.026	0.870
<i>devRate</i>	1.032*	0.000	1.033*	0.000	1.040*	0.000
<i>numLots</i>	1.003*	0.002	1.003*	0.007	1.004*	0.001
<i>reqOpenSpace</i>	1.487*	0.018	1.406*	0.039	1.439*	0.037
Opportunity costs						
<i>class1</i>	1.006	0.517	1.010	0.335	1.010	0.375
<i>class2</i>	1.003	0.133	1.004**	0.069	1.004*	0.073
<i>class3</i>	1.002	0.417	1.003	0.303	1.003	0.337
<i>class4</i>	1.004	0.331	1.004	0.314	1.004	0.341
<i>acres</i>	1.025*	0.001	1.025*	0.000	1.027*	0.000
<i>acres</i> ²	1.000*	0.016	1.000*	0.014	1.000*	0.000
<i>hasHouse</i>	0.174*	0.000	0.221*	0.000	0.138*	0.000
Construction costs						
<i>steep</i>	0.996	0.106	0.997	0.126	0.999	0.123
<i>forest</i>	1.003**	0.067	1.003	0.093	1.003**	0.084
<i>notRoadSuit</i>	0.987*	0.000	0.988*	0.000	0.987*	0.000
<i>notSepticSuit</i>	1.004	0.167	1.004	0.101	1.005	0.078
<i>sewerPlnd</i>	0.359*	0.000	0.329*	0.000	0.293*	0.000
<i>intRate</i>	0.918	0.154	0.802*	0.058	(a)	(a)
<i>Apfo</i>	0.693	0.127	0.860	0.549	0.899	0.678
Options variables						
<i>Easement</i>	0.499*	0.005	0.458*	0.003	0.424*	0.001
<i>Drift measure</i>	1.009	0.367	1.029**	0.051	1.028**	0.076
<i>Variance measure</i>	1.000	0.914	0.988	0.434	0.987	0.368
	η	1.34 (0.056)				

Note: Single asterisk (*) denotes 5%, double asterisks (**) 10% significance level.
(a) These variables do not vary over observations for each risk period, preventing estimation in the Cox model.

according to the number of developable lots, and allow different coefficients for the drift and variance for each quintile. Only the results with respect to the option-related variables appear in table 3, as only small changes in other coefficient estimates were observed in the expanded model.

In the “quintiles” models, the drift coefficients remain insignificant for all but 4 of the 15 estimated coefficients across the three specifications. The coefficients associated with the variance variables, however, are significant for four of the five quintiles in all three specifications and exhibit a clear increasing pattern with lot capacity. For parcels that can accommodate eleven lots or fewer (those in the 1st, 2nd, and 3rd quintiles), an increase in the variance decreases the hazard and delays development,

just as the theory suggests, but the delaying effect is smaller as the lot capacity increases.

For those parcels that can accommodate the largest number of lots, an increase in the variance actually appears to move development forward in time. There are several possible explanations. For one thing, the returns to scale implicit in housing development may differ substantially from the inherent assumptions of the theoretical real options models. Also, parcels with large development capacities are rare and declining in numbers in Howard County, so that fear of preemption may be hastening development. It is also possible that expectations of continually increasing regulation of large subdivisions may counteract the tendency for uncertainty to delay conversion of these parcels.

Table 3. Selected Hazard Model Estimates from Expanded Model

Variable	Number of lots	Weibull		PW Exponential		Cox – Exact	
		Log(L) = -978.931		Log(L) = -1000.570		Log(L) = -2530.146	
		Ratio	$p > z $	Ratio	$p > z $	Ratio	$p > z $
Drift_0–20%ile	3 – 5	1.053	0.165	1.078**	0.056	1.070**	0.058
Drift_21–40%ile	6 – 7	0.989	0.782	1.016	0.715	1.015	0.723
Drift_41–60%ile	8 – 11	0.977	0.468	1.002	0.942	1.001	0.896
Drift_61–80%ile	12–22	1.023	0.170	1.041*	0.028	1.042*	0.027
Drift_81–100%ile	23–859	1.006	0.704	1.017	0.351	1.014	0.352
Var_0–20%ile	3 – 5	0.942*	0.006	0.930*	0.002	0.931*	0.001
Var_21–40%ile	6 – 7	0.951*	0.029	0.940*	0.008	0.940*	0.003
Var_41–60%ile	8 – 11	0.964**	0.064	0.953*	0.021	0.954*	0.020
Var_61–80%ile	12–22	1.010	0.522	0.998	0.908	1.001	0.951
Var_81–100%ile	23–859	1.050*	0.001	1.037*	0.020	1.044*	0.014
Easement		0.385*	0.000	0.345*	0.000	0.331*	0.000

Note: Single asterisk (*) represents 5%, double asterisks (**) represent 10% significance level.

Easement Option Results

By far the most dramatic results are those attached to the easement option. We find robust evidence that the addition of a preservation option significantly delays the development decision for qualified parcels, lowering the hazard rate by at least 50% (table 2). In each specification the coefficient on the easement variable is significant at the 99% confidence level. With the expanded model, in which the drift and variance are allowed to vary with parcel size, the effect of qualification for an easement remains significant, and is larger in size, reducing the hazard of development by at least 63% (table 3).

These are quite large estimated effects and there is a reason why they may be overestimates. As we explained earlier, parcels that preserve during the study period pose a practical problem. These parcels are legitimately at risk for development until their owners choose to preserve them. However, once their easements have been sold, they are no longer at risk. This suggests that retaining them in the risk set until the date of preservation and then dropping them from the analysis is the appropriate treatment. An alternative treatment is to leave the preserved parcels in the risk set and include a dummy variable for the preservation state, in recognition of the fact that it is *technically* feasible for them to be developed although at a very large cost to the owner. Some programs contain a provision allowing landowners to buy back the easement after a specified period; the hardship conditions that must be met are sufficiently stringent that no landowner has yet attempted it. Thus, in principle the preservation option might be consid-

ered reversible, but only at a very high cost. The qualitative results reported in table 2 do not change appreciably under this treatment of preserved parcels, but the estimated effect of the preservation option is smaller. The preservation option is estimated to decrease the hazard rate by about 46% and is statistically significant at a 99% confidence level.

Either treatment has the drawback of potentially affecting our interpretation of the easement variable coefficient, however. Suppose that a landowner whose parcel is eligible for preservation makes the decision to preserve. The actual easement sale may occur after the time the parcel would otherwise have been expected to develop. This may occur if the enrollment process is time-consuming or if county budgetary limitations delay the sale. During this interim period, the parcel remains in the risk set beyond the time it would have been predicted to develop, and does not drop out of the risk set until the time at which it ultimately preserves. These “pending” preservations could contribute econometrically to the apparent delay in development resulting from easement eligibility.

In order to make the claim that simply having the option to preserve is sufficient to delay development, even if a parcel is not ultimately preserved, we need to rule out the possibility that the delaying effect of the easement dummy is due entirely to delay in intended preservations. One simple way to address this question is to reestimate the models eliminating the parcels that preserve during the study period. This procedure is less desirable as it discards useful information, but it helps us establish more conclusively that we have identified a real effect. Using this subset of 1,627 parcels,

the effect of the preservation option remains significant at the 95% level in all three model specifications although it is reduced in magnitude: the preservation option is estimated to lower the hazard of development by at least 39%. Further subsetting the data by excluding all censored parcels, leaving just the 383 parcels that developed during the study period, is even more wasteful of data, but produces additional support for the existence of the preservation option effect. With this circumscribed data set, the easement option is still found to lower the estimated hazard of development by 36% in the semiparametric specifications.⁷

Using the estimated coefficients from table 2, it is possible to predict the median time until conversion (median duration time) for the censored parcels in the data set. We did so for each of four-size categories of parcels (starting at 25 acres—the minimum size of preservation eligible parcels), basing the calculations on the estimated results from the Weibull model.⁸ The predicted median durations are significantly longer for parcels with a preservation option, with the median duration ranging from about twenty to forty-six years (depending on parcel size) for preservation-eligible parcels, as contrasted with median durations of about fourteen to forty-one years for preservation-ineligible parcels. These results suggest that the mere existence of the preservation option extends the time until conversion by an average of almost six years, a delay of 12–43% depending on size. Predicted duration rates decline nonlinearly as parcels increase in size.

Summary

This study produces empirical estimates of real options theory phenomenon in a land use context. Of principal interest is an empirical test of whether the presence of an alternative land use option—specifically, an option to preserve farmland in a PDR program—delays development decisions. The policy significance of preservation options in particular is that if these options do induce delays for parcels that ultimately convert, PDR programs may

be providing more benefits than those simply associated with the parcels that are enrolled in the programs. We use a duration modeling approach, which explicitly accounts for the probability a given parcel will convert in the next period is conditional on the fact it did not convert in any previous period. We find statistically significant evidence that the option to sell a PDR easement decreases the rate of development. This finding is in line with real options theory predictions: the addition of an option to the choice set is expected to increase the value of waiting.

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 did not find strong support for the contention that increasing drift in returns delays development of parcels. At least two data-related reasons may account for the limited results. 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 trends, 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 such an option was predicted to extend the median duration (the delay until conversion) by about six years for parcels that had not yet exercised a development or preservation option by the end of our study period.

While amenity benefits are provided by both farmland that is enrolled in PDR programs and farmland whose conversion is delayed due to the easement option, this study suggests the amenity benefits generated by postponing the loss of farmland may be limited because the induced delay is only a few years. However, if our qualitative results turn out to be generalizable to other regions—especially those with lower overall development pressures—then

⁷ The coefficient on the easement dummy under the parametric Weibull specification remains correct in sign but loses significance. However, a nonparametric log rank test using just the set of parcels that developed reveals the distribution of hazard rates are significantly different between easement eligible and noneasement eligible parcels. This suggests the easement results are not solely arising from parcels waiting to preserve.

⁸ We chose this specification over others based on the Akaike Information Criteria.

delays due to the additional option may be longer and “postponement” benefits considerable. Even in fast growing areas, delays may be beneficial if the public sector is struggling to keep up with rising infrastructure demands. If a few years’ delay gives the local government time to institute stricter growth control measures, the effect of the delay due to the existence of an alternative option may be even more long term. Also, the finding that PDR programs delay conversion suggests that relaxing eligibility constraints could provide more amenity and planning benefits by delaying the conversion of more parcels. These potential benefits could be tempered, though, if development on parcels *not* eligible for preservation is consequently hastened.

Whether our results are generalizable to other PDR programs may depend, in part, on how similar important program features are to Howard County’s program. One important feature is how much the PDR program pays a landowner for an easement, relative to what the easement is worth. Howard County’s use of a static formula to determine the easement price may make its 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 (Geltner, Riddiough, and Stojanovic 1996). Increased transaction costs associated with the preservation option may also work in the reverse direction. 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.

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