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Elena G. Irwin

ABSTRACT. *The marginal values of different open space attributes are tested using a hedonic pricing model with residential sales data from central Maryland. The identification problems that arise due to endogenous land use spillovers and unobserved spatial correlation are addressed using instrumental variables estimation with a randomly drawn subset of the data that omits nearest neighbors. Results show a premium associated with permanently preserved open space relative to developable agricultural and forested lands and support the hypothesis that open space is most valued for providing an absence of development, rather than for providing a particular bundle of open space amenities.* (JEL R52)

I. INTRODUCTION

The decision to preserve open space is often justified based on the value of the natural amenities associated with the land, for example, the biodiversity, wildlife habitat, or scenic views provided by the land. However, rather than being valued for what it is, some evidence suggests that open space may more often be valued most for what it is not—that is, for not being development. For example, Halstead (1984) and Beasley, Workman, and Williams (1986) estimate that households' willingness to pay to preserve an acre of average quality farmland increases from \$50 to \$150 per household when the replacement for agriculture is hypothesized to be high density rather than low density development. Other contingent valuation studies have also demonstrated a positive willingness-to-pay for farmland preservation as a means of preventing development (Bergstrom, Dillman, and Stoll, 1985; Krieger, 1999; Bower and Didychuk, 1994). As summarized by Heimlich and Anderson (2001), these studies estimate a wide range of willingness-to-pay estimates, from annual values of \$0.21–\$49.80 for 1,000 acres of preserved farmland that is expressly prevented from being developed. However, while these estimates are useful for public policies regarding open space preservation, they do not clarify how individuals trade-off particular types of open space, for example, cropland vs. forest, nor do they shed light on the extent to which the particular attributes of open space may be secondary to the absence of development that preserving open space provides.

An alternative approach to assessing the value of open space is the hedonic pricing method. Evidence of the value of open space using this approach has provided some estimates regarding the marginal values of different types of open space, but results from these studies are mixed. Tyrävinen and Miettinen (2000) conduct a careful study of the spillover effects of the various attributes associated with urban forests on housing prices in a semi-rural area in Finland. They find that the distance to the nearest small area of forest has a negative and significant effect, and that the presence of a forest view from the housing unit has a positive influence. Other open space variables are not found to be significant, including the relative amount of forested area within the housing neighborhood and the distance to the nearest large forested area. Garrod and Willis (1992) also focus on the value of forests by testing whether different types of tree stocks affect neighboring housing prices. They find that deciduous trees located within the same one-kilometer grid as residential homes significantly increase house prices, but that spruce conifers significantly decrease house prices. Alternatively, Geoghegan, Wainger, and Bockstael (1997) examine the aggregate effect of sur-

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rounding agricultural and forested lands on the value of residential exurban land in a central Maryland region. Differing open space effects are found, depending on the size of the neighborhood considered: within a tenth of a kilometer radius, the proportion of open space positively impacts land values, but within a one-kilometer buffer negatively influences land prices.

Other studies have considered whether there are differing effects from open space that is ''developable'' vs. open space that is in some way preserved. Cheshire and Sheppard (1995) estimate the effects of publicly and privately held open space on residential property values using separate datasets from two medium-sized towns in England. They find different results that depend on the amount of open space amenities in the two towns. Only if the amount of either public or private open space is sufficiently scarce in the towns is a positive and significant effect found on residential property values. Irwin and Bockstael (2001a) consider the effects of privately held open space that is developable vs. privately or publicly held open space that is permanently preserved and find that each type of open space generates positive and significant spillover effects. However, they do not test whether the estimated effects are significantly different from each other. Geoghegan (2002) considers the effects of developable vs. permanent open space and finds that permanently preserved open space is three times more valuable than open space that could be developed at some point in the future.

The differences in these results regarding the marginal values of different types of open space demonstrate that open space itself is a heterogeneous good. Open space may be distinguished by land use, land cover, ownership type, development potential, and geographic location, each of which may be valued differentially. However, because studies have tended to focus on a particular type of open space or have aggregated different types of open space into one aggregate category, much less evidence exists regarding the relative values of the various attributes that are associated with open space.

Using residential sales data from an exurban region in central Maryland, this article employs a hedonic pricing model to test whether different types of open space generate significantly different spillover effects. Open space is distinguished first by whether the land is preserved or is developable, and second by land ownership (privately vs. publicly held preserved open space) and land use type (cropland, pasture, and forests that are developable). In doing so, the goals are to explore whether preserved open space carries a premium with it and whether the various landscape amenities that are associated with different open space land uses have differing marginal values. Based on these results, we hope to draw conclusions regarding the extent to which open space is valued for its particular attributes vs. for simply not being development.

The rest of the article is organized as follows. First, we briefly discuss the hedonic pricing model and its use as a means of estimating the marginal value of landscape attributes. This is followed by a discussion of the identification problems that can arise in estimating land use spillovers with a hedonic-pricing model, as outlined in Irwin and Bockstael (2001a). These are addressed by using an instrumental variables estimation approach with a randomly drawn subset of the data that eliminates nearest neighbors to control for the bias introduced by endogenous variables and unobserved spatial heterogeneity, as well as the inefficiency caused by spatial error autocorrelation. Results from the model are presented and the robustness of these results to the spatial sampling routine are explored. Finally, the marginal values for various types of open space lands are derived and the implications of the findings for open space preservation policies are discussed in the concluding sections.

II. HEDONIC MODELS AND IDENTIFYING LAND USE SPILLOVERS

Hedonic pricing models offer a means to estimate the marginal implicit prices of characteristics associated with a differentiated market good, such as housing. The hedonic price function, which posits price as a function of the quantities of a good's attributes, arises through the interactions of many buyers and sellers in the market. As a result, it describes the locus of equilibrium points between buyers and sellers in the market. The marginal implicit price of any of the good's attributes is found by differentiating the hedonic price function with respect to the attribute. Evaluated at an individual's optimal choice, this implicit price represents the individual's marginal willingness-to-pay for the attribute. Because housing is spatially immobile, the values of location amenities, such as open space spillovers, are capitalized in the sales price of the home. Although this approach is limited because it fails to capture any value from open space that does not accrue to nearby residents, it provides a partial estimate of open space benefits, which is useful for evaluating the trade-offs associated with open space preservation.

For the purposes of this paper, we specify the hedonic residential pricing model as:

$$P_i = f(\mathbf{H}_i, \mathbf{N}_i, \mathbf{L}_i; \boldsymbol{\beta}, \boldsymbol{\lambda}, \boldsymbol{\delta}), \quad [1]$$

where P_i is the residential sales price of the i^{th} property, $\mathbf{H}_i$ is a vector of structural characteristics associated with the house, $\mathbf{N}_i$ is a vector of neighborhood/locational variables, $\mathbf{L}_i$ is a vector of neighborhood land use variables, each of which measures the proportion of the surrounding land that is in a particular land use, and $\boldsymbol{\beta}$, $\boldsymbol{\lambda}$, and $\boldsymbol{\delta}$ are the respective parameter vectors to be estimated.

A variety of econometric issues arise in estimating hedonic models, including questions of functional form, extent of the housing market, and problems associated with multicollinearity and spatial correlation. An issue that is specific to the estimation of land use spillovers using hedonic pricing models is the identification of these effects given the potential endogeneity and spatial correlation of the neighborhood land use variables. The potential endogeneity of open space variables arises when a particular type of open space land is included in the analysis—specifically, open space that has the development rights intact and that can be converted to a residential use at any point in the future.[1] As discussed by Irwin and Bockstael (2001a), if open space can be developed as residential land use, then it is part of the market for residential land and subject to the same economic forces that determine a location's residential value. This implies that variables measuring the influence of this particular type of open space on neighboring residential property values will be endogenous in a hedonic pricing model. As a result, identification problems arise that will bias the open space coefficients.

Specifically, Irwin and Bockstael argue that two identification problems arise in this context. The first is the standard type of econometric identification problem due to endogenous explanatory variables. Consider two neighboring parcels, i and j, both of which consist of a varying amount of residential land use and/or open space that may be developed in the future. The amount of residential or open space land use on parcel j is influenced by its value as a residential location, which, because of the spillover effect, is a function of land use spillovers from parcel i. In turn, the amount of residential development and open space on parcel i is determined by its residential property value, which is a function of the land use spillovers from parcel j. Therefore, the residential value of parcel i is a function of the residential value of parcel j and the measure of surrounding open space around parcel i, which is a function of parcel j's residential value, is endogenous.

The second problem arises because spatial error autocorrelation is likely in hedonic models. In the standard case, in which the explanatory variables are exogenous, this leads to an inefficiency problem: the standard errors of the estimates are biased, but the estimated coefficients themselves are not (Anselin 1988). However, if the open space

[1] Such a situation—in which surrounding open space can be converted to residential use—is typical of many exurban and rural places in the United States in which the default zoning for undeveloped land is residential, but it is not ubiquitous. For example, European countries, by and large, have different institutional arrangements that either prohibit or greatly constrain conversion of such land.

variables are endogenous, then they will be spatially correlated with the error term, which creates a second source of bias. In addition, to the extent that the spatial error correlation is time invariant, this creates potential problems for measuring the effects of spillovers from existing development on neighboring land. Consider again the neighboring parcels i and j and suppose that parcel j was converted in some past period to a residential land use. Assume that its current land use is a function of its past residential value, $P_{j,t-\tau}$, where $t - \tau$ represents the period in which parcel j was converted to a residential use. If a portion of the unobserved spatial correlation that influences a parcel's residential value is also time invariant, then this implies a pattern of space-time autocorrelation in which ε_{it} and $\varepsilon_{j,t-\tau}$ are positively spatially correlated. Therefore, to the extent that parcel j's current land use is a function of its past residential value in period $t - \tau$, this time invariant, spatial correlation will cause correlation between the measure of neighboring development around parcel i in period t (which includes parcel j's residential development) and the error term in the hedonic pricing model, ε_{it}. As a result, the estimates associated with these land use spillovers will be biased.

Other econometric issues that arise in estimating hedonic pricing models include model specification, functional form, extent of the market, and multicollinearity. Ideally, model specification and functional form are guided by theoretical considerations. However, other than theoretical guidance regarding the expected signs of certain coefficients, there is little other guidance regarding model specification or restrictions to the functional form. Model specification is often dictated by data availability and *a priori* beliefs about the types of location and structural amenities that matter to households. A number of questions that are specific to the specification of the neighborhood land use variables arise, including the relevant size of the neighborhood and the degree to which individuals' perceptions of land use spillovers correspond to the distinction of the land use categories used in the model.

Because of the lack of theoretical guidance regarding the choice of a functional form, this is typically informed by empirical evidence. A common approach is to compare goodness-of-fit criterion from alternative functional forms (e.g., log-log vs. semilog) and choose the best fitting model in this way. Alternatively, a Box-Cox transformation can be used to generalize the model, in which one or more additional parameters are introduced whose values specify the functional form of the dependent and independent variables.[2] These parameters are treated as unknown and are estimated along with the other parameters of the model. Given the unrestricted estimated values, a likelihood ratio test can be used to determine whether a restricted model, in which a particular functional form is assumed, imposes a significant restriction on the parameter values. While the Box-Cox transformation may be performed using linear, quadratic, or other functional forms, there is some evidence in the literature that, when omitted variables are a potential problem, a linear version of the Box-Cox transformation is generally the most robust (Cropper, Deck, and McConnell 1988).

The question of the geographic extent of the housing market often arises in estimating hedonic pricing models. Some have argued that regional housing markets are more accurately represented as an aggregate of smaller, distinct housing markets, implying that separate hedonic price functions should be estimated for separate geographic areas of the region. This approach relies on the assumptions that either the structure of demand or supply is different across the segments and that there is not significant overlap (in terms of buyers and sellers) across the market segments (Freeman 1993). If these conditions hold, then estimation of separate hedonic functions that correspond to the separate market segments is warranted.

Lastly, multicollinearity is often a problem in estimating hedonic models of residential housing values, in part because of the evolutionary nature of the urban spatial struc-

[2] The Box-Cox transformation of a variable is: $X^\lambda = (X^\lambda - 1)/\lambda$. For $\lambda = 1$, this is a simple linear function. As λ approaches zero, it becomes a log function (by L'Hôpital's Rule).

ture process. For example, to the extent that households with similar preferences are attracted to the same types of neighborhoods, a number of physical and socio-demographic attributes of the neighborhood and structural attributes of the houses are likely to be correlated, that is, neighborhoods with larger houses will, on average, have higher income households residing in them and will be located within a certain range outside the central city. Multicollinearity, if present, can lead to low significance levels due to high standard errors and large changes in parameter estimates given a small change in the data or model specification.

DATA

The study area is comprised of suburban and exurban counties within a central Maryland region that belong to the Washington, D.C.-Baltimore metropolitan area. These counties include Anne Arundel and Howard counties, both of which contain significant suburban and some urban population, as well as Calvert and Charles counties, both of which are largely exurban in nature. The data consist of 55,799 arms-length, single transactions of owner-occupied residential properties that occurred within these counties between January 1995 and December 1999. The data are from the Maryland Office of Assessment and Taxation and are made available in geocoded format by the Maryland Office of Planning through Maryland Property View, a GIS data product. Additional spatial variables were generated using ArcInfo GIS software.

In distinguishing types of open space, different classifications are possible, that is, by land use, land cover, ownership type, development potential, or geographic location. Ideally, the classification would reflect individuals' own perceptions of different types of open space. However, this information is not at hand and therefore the distinctions are drawn based on maintained assumptions regarding individuals' perceptions of neighboring open space. We surmise that people distinguish land first by whether it is in a preserved state vs. being developable, and second by its land ownership and land use. Using this approach, the first distinction is between privately owned open space with development rights intact, that is, land that could be developed at anytime vs. land that has been permanently preserved in some way. Permanently preserved lands are then further distinguished based on whether they are privately owned land whose development rights have been sold or land that is publicly held. Public lands are further distinguished by whether they are military land or not; and privately owned lands that are developable are further distinguished based on their defining land use/landscape attribute: cropland, pasture, or forest. This categorization of open space yields six different measures of surrounding open space: (1) cropland that is privately owned (CROP); (2) pasturelands that are privately owned (PAST); (3) forested lands that are privately owned (FOREST); (4) privately owned land that is protected from development, including agricultural easements and privately owned conservation areas (CONSV); (5) non-military open space land owned by the federal, state, or county governments (PUBLIC); and (6) military land that is in open space owned by the federal government (MILIT). Because Anne Arundel County contains large areas of military land that are in open space, including what is by far the largest contiguous area of open space in the study area (Fort Meade), and very few non-military open space areas that are publicly owned, we interact the public land variables with the Anne Arundel County dummy variable to control for these potential differences within Anne Arundel County (AAPUBLIC and AAMILIT).

Each of these values is measured in proportionate terms and indicates the percentage of the total land area within a specified neighborhood that is classified as a particular type of open space. The specification of the neighborhood extent is largely an empirical question and can, in part, be determined by initial exploratory data analysis of the spatial pattern. Here, visual inspection of the distribution of residential properties relative to the pattern of surrounding land uses in the study area suggests that land uses within an immediate vicinity of residential parcels are contained within a 400-meter radius. Based on this, the neighborhood is specified as a buffer around each parcel centroid defined by a 400-meter radius.

In addition to the neighboring open space land use measures, three additional measures of land use spillovers are included to capture the externality effects of neighboring development: the proportion of neighboring land that is in low density residential land use (LOWRES); medium and high density residential development (MEDHRES); and commercial or industrial land use (COMIND). Lastly, we include a ''catch-all'' measure to control for the net effect of all other land uses that are within the specified neighborhood of each parcel (OTHER). As a result, because all land uses within a parcel's neighborhood are measured and are defined as proportions variables, their sum is equal to one. The model is estimated, therefore, by normalizing on a base land use that is dropped from the model. The interpretation of the parameters is relative to this base land use and, because the amount of land within the neighborhood is fixed, the interpretation is in terms of the marginal value of a change in the proportion of land that is converted from the base land use to another. For example, if cropland is dropped from the model, then the estimated forest coefficient would give the change in value to a residential property given a one percent increase in the amount of neighboring land that is converted from cropland to forest.

To control for other variations across location, several additional location-specific variables are included. The location of parcels relative to major urban centers is likely to matter. Measures of the distance via the major roads network to the two major centers in the study area, Washington, D.C. (DISTDC) and Baltimore, Md (DISTBA), are included. Because the study region contains a large, international airport (Baltimore Washington International), a separate variable for aircraft noise is included to distinguish the potential disamenity of living nearby. This variable is a dummy variable (AIRPORT) that takes on the value of one if a residential property is located within one mile of the airport and zero otherwise. Several socioeconomic variables, taken from the 1990 *U.S. Census of Population* and measured at the block group level, are included as measures of neighborhood quality: median household income (MHHINC), population density (POPDEN), and the percentage of the neighborhood population that is African-American (BLPOP). *A priori*, we expect that residential sales price will decrease with population density and the percentage of the population that is African-American and increase with median household income. Most public services, including public schools, are provided on a county level in Maryland. To control for differences in these services, we include county dummies for three of the four counties in the study area, Calvert (CA), Charles (CH), and Howard (HO) counties, and omit the dummy variable for Anne Arundel County.

Because we are primarily interested in the value of the open space amenities provided by the landscape, these values should be reflected in the land price itself. However, the estimation is with data on market transactions of houses and therefore the inclusion of structural characteristics is necessary to control for variations across housing stock. Several different structural attributes are included in order to control for this. These include an index variable that rates the grade of the dwelling unit on a scale of 0–9, with 9 being the highest grade (DWGRADE); a dummy variable indicating whether the dwelling is a detached unit (DWTYPE); number of full baths (BATHS_FU); number of half baths (BATHS_HA); the square feet of the structure (AREA); the footprint of the house (FTPRNT); the age of the house (AGE); and the year of the sale (YRSALE). Lastly, lot size is hypothesized to influence the residential value of a property (LSIZE).

EMPIRICAL RESULTS

Drawing upon the evidence discussed earlier that choosing a simpler functional form is generally more robust to situations in which omitted variables may be a problem, we compare the adjusted R^2 values from a log-log, semilog, and a simple linear form of the model. In the log-log model, all the right hand side variables are expressed in terms of logs with the exception of dummy variables

TABLE 1
OLS ESTIMATION RESULTS

Dependent Variable: log(price)
Number of Observations: 55,799
R^2: 0.7127
Adj R^2: 0.7126

Variable	Parameter Estimate	Standard Error	t-Value
Intercept	3.89437*	0.10776	36.14
DWGRADE	0.15671*	0.00242	64.72
DWTYPE	0.17258*	0.00392	43.98
BATHS_FU	0.07107*	0.00219	32.43
BATHS_HA	0.05302*	0.00237	22.39
FTPRNT	0.10106*	0.00555	18.21
AREA	0.34139*	0.00499	68.45
LSIZE	0.02337*	0.00153	15.32
AGE	−0.02158*	0.00102	−21.23
YRSALE	0.02219*	0.00075808	29.27
DISTBA	0.06991*	0.00317	22.08
BWI	−0.01090*	0.00258	−4.22
DISTDC	−0.07736*	0.00465	−16.62
MHHINC	0.18020*	0.00521	34.6
POPDEN	−0.01548*	0.0009972	−15.52
BLPOP	−0.06475*	0.00905	−7.15
CA	−0.27684*	0.00596	−46.42
CH	−0.22990*	0.00567	−40.58
HO	−0.10029*	0.00333	−30.08
LOWRES	0.06271*	0.01332	4.71
COMIND	−0.07968*	0.0173	−4.61
MEDHRES	−0.03378**	0.01213	−2.78
CROP	0.01192	0.0152	0.78
FOREST	0.02577^+	0.01304	1.98
CONSV	0.27483*	0.0737	3.73
PUBLIC	0.07764*	0.02391	3.25
MILIT	0.21534	0.14009	1.54
OTHER	0.23814*	0.01401	17
AAPUBLIC	−0.15521*	0.02774	−5.6
AAMILIT	0.52320*	0.14892	3.51

*, **, and $^+$ indicate significance at the 0.001, 0.005, and 0.05 levels respectively.

and proportions variables that vary between zero and one. Results show a clear dominance of the log-log and semilog specifications over the linear model and a slight preference for the log-log model relative to the semilog model.[3]

Table 1 presents the results from the OLS estimation of the log-log model. The model is estimated for the entire study region, thus with the assumption that the study region comprises one regional housing market.[4] All of the housing attributes are significant at the .0001 level and are of the expected sign. Residential housing price is increasing in the grade and type of housing unit and increasing at a decreasing rate in the size of the house, as measured by the footprint and number of square feet, as well as the size of the lot. In addition, sales price is decreasing at a de-

[3] Adjusted R^2 for the log-log, semilog, and linear models are 0.711, 0.694, and 0.031, respectively.

[4] In reality, the market is likely comprised of several local, but overlapping markets for housing. However, because the study region is contained within the Washington D.C.-Baltimore metropolitan area, significant overlap among these smaller housing markets is expected. In addition, no physical features, such as a major river or mountain range, exist to separate these areas and all counties contain major roads that connect them to the major interstates that serve the region. For these reasons, the housing market is believed to be better represented as one regional market rather than several geographically distinct markets.

creasing rate with the age of the house. All of the locational variables are significant at the .0001 level and most are of the expected sign as well. The value of a residential property is found to be increasing at a decreasing rate in the median income level of households living within the block group in which the property is located and decreasing at a decreasing rate in the population density of the same area. Price is decreasing in the proportion of the local population that is African-American. All three of the county dummies included in the model are negative and significant, reflecting the fact that these counties are less desirable as a residential location than Anne Arundel County, *ceteris paribus*.

Consistent with the basic premise of the urban bid rent model, residential prices are decreasing (at a decreasing rate) with distance from the major urban center of Washington, D.C. Counter to expectations, a positive bid-rent gradient is found for Baltimore, the other major urban center in the study region. While it is true that the land just south of Baltimore contains many pockets of industrial activity and that average housing values are lower there, this result is puzzling since the model controls for the disamenities from surrounding industrial and commercial land, as well as proximity to the airport that is located approximately ten miles south of the city. Given this, it is likely that there are additional disamenities associated with this area that are not controlled for by the model. An alternative, but less likely, interpretation is that other employment centers and destination sites within this region that are not included in the model dominate Baltimore, implying that Baltimore's urban spatial structure is non-monocentric.

Of primary interest are the measures of land use spillovers and in particular, the open space measures. The land use variables are normalized to the developable pastureland variable, so that the results indicate the marginal spillover effect of a neighboring land use relative to pasture. It is not obvious *a priori* which types of open space would be expected to confer greater value. To the extent that preserved open space is perceived as being guaranteed open space into the future, this reduction in uncertainty may carry a premium with it. However, public open spaces are often destination sites for people from outside the local area and may also generate a nuisance spillover if they lead to less privacy and greater congestion than privately held open space. Among privately held open space, pasturelands might be expected to confer greater value than either crops or forests because these lands generally offer more scenic views.

The results show that both the privately owned conservation lands and public, non-military open space (CONSV and PUBLIC) have a positive and significant effect on the value of neighboring residential properties relative to developable pastureland. In addition, the coefficient associated with developable forests (FOREST) is positive and significant at the .04 level. However, the coefficient associated with privately owned cropland that is developable (CROP) is positive, but not significant relative to cropland. Military land (MILIT) is not found to have a significant effect. Other land use spillovers from development are found to have a significant effect: higher density residential (MEDHRES) and commercial/industrial land uses (COMIND) are both estimated to have a negative effect relative to spillovers from neighboring pasture and low density residential land use (LOWDRES) is estimated to have a positive influence on neighboring residential housing values relative to neighboring pastureland. The public land variables that were interacted with the Anne Arundel County variable reveal that these effects operate differently in this county. The separate effect of military lands within Anne Arundel County (AAMILIT) is found to have a positive and significant effect on surrounding residential property values, reflecting the positive amenity that the open space spillovers from Fort Meade provide. Lastly, the separate effect of publicly held, non-military open space within Anne Arundel County (AAPUBLIC) is found to have a negative and significant effect on neighboring residential property values.

As discussed earlier, multicollinearity is often a problem in estimating hedonic models. However, diagnostic tests indicate that it is not a significant problem here.[5] The identification issues relating to endogenous variables and unobserved spatial correlation, however, do prove to be a problem. These problems arise because the estimated OLS coefficients associated with neighboring land uses that are a part of the residential land market may be biased due to an endogeneity problem. If omitted variables are spatially correlated, this problem is further complicated by the resulting correlation between the spatially correlated error and land use variables. As discussed earlier, this second problem also causes bias estimates and, to the extent that the spatial correlation is also time invariant, is potentially a problem with land use variables that are a function of their past residential values. We would certainly expect that existing residential development to be a function of the parcels' residential value at the time of conversion and, to the extent that commercial development is a function of neighboring residential development, it is reasonable to expect that this land use would be a function of parcels' past residential values. A Lagrange Multiplier test for spatial error autocorrelation indicates that the errors from the OLS estimation are significantly positively spatially correlated.[6] A joint Hausman specification test of the coefficients associated with the developable open space measures (CROP and FOREST) indicates that the OLS estimates are significantly different from instrumental variables (IV) estimates.[7] Furthermore, this same test with these developable open space measures and developed land variables (LDNRES, HDNRES, and COMIND) indicates that the OLS estimates associated with all five variables are jointly significantly different from the instrumental variables (IV) estimates.[8] Therefore the IV estimation is performed using exogenous features of the landscape as instruments for the five variables listed above that measure the spillovers from privately held and developable land uses and existing development. The instruments that are included are: (1) a dummy variable indicating the steepness of a parcel's slope. This variable takes on the value of 1 for parcels that have steep slopes (more than 15%); (2) a dummy variable indicating the drainage potential of the soils that takes on a value of 1 if the parcel has poorly draining soils and 0 otherwise; (3) a dummy variable that indicates the quality of the soils for agriculture, which equals 1 for parcels that are currently employed in agricultural activity and have prime agricultural soils,[9] and 0 otherwise; and (4) distance from the two urban centers, Washington, D.C., and Baltimore (in log form). These instruments are believed to be exogenous to the residential housing market, and therefore uncorrelated with the error term, but correlated with the spatial pattern of open space and development. For example, steep slopes and poor soil drainage are expected to increase the costs of converting a parcel and therefore the dummy variables that indicate these features should be positively correlated with the pattern of developable open space and negatively correlated with the development pattern. Likewise, high quality agricultural soils should be positively correlated with agricultural land and negatively correlated with development, whereas distance to urban centers is expected to be negatively correlated with open space and positively correlated with development.

While the IV estimation controls for the bias introduced by the endogenous variables and unobserved spatial correlation, it does not correct for the inefficiency of the esti-

[5] The condition number of the scaled data matrix is 9.8. This number, which is the square root ratio of the largest and smallest eigenvalues associated with the scaled data matrix, indicates multicollinearity problems when in excess of 20 (Greene 2000).

[6] The value of the Lagrange Multiplier statistic is 17.611. This statistic is distributed chi-squared with one degree of freedom. It is significant at the 0.01 level.

[7] The value of the joint Hausman test statistic is 37.79. This statistic is distributed chi-squared with five degrees of freedom and is significant at the 0.001 level.

[8] The value of the joint Hausman test statistic for the case in which all five OLS parameter estimates are hypothesized to be biased is 1565.08, which clearly indicates that the IV and OLS estimates are significantly different.

[9] The soil and slope attributes are defined according to the Maryland Department of State Planning, Natural Soil Groups of Maryland, Technical Series Publication 199 (December, 1973).

mates caused by the remaining spatial error correlation. As such, hypothesis testing is suspect since positive spatial error correlation will bias the standard errors upwards. To control for this problem, a randomly drawn subset of the data is created in which nearest neighbor observations are dropped from the dataset. This approach, which is further detailed in Haining (1993), is often used to control for spatial error correlation when the standard method of specifying a spatial weights matrix and estimating a spatial error model is not possible or not appropriate.[10] In our case, a spatial weights matrix approach is not appropriate since it is not possible to separate the endogenous spatial spillover effects from the spatial error, given that they are both positively correlated.[11] Nonetheless, this approach is limited by the maintained assumption of how nearest neighbors are defined. In what follows, nearest neighbors are first defined as parcels that are within 100 meters of each other and then, to test the robustness of this assumption, we systematically alter the definition of nearest neighbors to be parcels within 200, 400, and 600 meters of each other.

Results from the instrumental variables estimation using the 100-meter buffer definition of nearest neighbors are reported in Table 2. With the exception of the airport dummy, the estimated coefficients associated with housing, locational and neighborhood socioeconomic characteristics remain unchanged in sign and significance. In addition, the estimated influence of privately owned conserved open space and publicly owned non-military lands (CONSV and PUBLIC) remains positive and significant and relatively constant. Both are found to generate additional benefits to surrounding residential values relative to developable pastureland. The coefficient associated with surrounding cropland (CROP) is positive, but not significant, indicating that the spillovers from surrounding pastureland are neither significantly greater nor lesser than those that are associated with cropland. On the other hand, the coefficient associated with surrounding forested lands (FOREST), is negative and significantly different from zero, indicating that the value of neighboring residential properties decreases as more land is taken out of pastures and placed into forests. The estimated magnitudes of the negative spillovers from surrounding commercial and industrial land use (COMIND) and from higher density residential land (MEDHRES) have increased. Lastly, the coefficient associated with surrounding low density residential land (LOWRES) is now estimated to be negative and significant at the .05 level, indicating that the value of a residential property decreases with a marginal change in the surrounding landscape from pasture to low density residential use.

THE VALUE OF OPEN SPACE PRESERVATION

To evaluate the marginal values of these open space effects, the first stage estimates from the hedonic pricing model that are reported in Table 2 and the mean values of all explanatory variables are used to calculate the change in the mean property's predicted price given a change in the neighboring landscape from one acre of pastureland to another land use. This approach is distinguished from use of the second-stage estimates, which are derived from the underlying inverse demand functions, to evaluate welfare changes. It relies on the assumption that the change being considered is small relative to the regional housing market, since it assumes that the hedonic function itself is stable. Given that the regional housing market considered in this study is approximately 1,350 square miles and that land use spillovers affect only a very localized area around them, use of the first stage estimates in this case is reasonable.

Using this method, we find that the conversion of one acre of developable pastureland to privately owned conservation land within a parcel's neighborhood increases the residential value of the mean

[10] See Nelson and Hellerstein 1997, for an example applied to land use change.

[11] This identification problem is well-established in the empirical literature on social interactions (Manski 1993; Brock and Durlauf 2001), but has not received as much attention in the spatial econometrics literature. For further discussion of this issue within the context of land use spillovers, see Irwin and Bockstael 2001b.

TABLE 2
IV ESTIMATION RESULTS (100-METER SUBSET)

Dependent Variable: log(price)
Number of Observations: 41,201
R^2: 0.5398
Adj R^2: 0.5394

Variable	Parameter Estimate	Standard Error	t-Value
Intercept	5.444006*	0.339172	16.05
DWGRADE	0.168764*	0.004057	41.6
DWTYPE	0.171619*	0.007575	22.66
BATHS_FU	0.071506*	0.003741	19.12
BATHS_HA	0.055979*	0.004129	13.56
FTPRNT	0.084669*	0.009504	8.91
AREA	0.343859*	0.008645	39.78
LSIZE	0.038042*	0.002703	14.07
AGE	−0.02003*	0.001752	−11.44
YRSALE	0.025194*	0.001308	19.27
DISTBA	0.096822*	0.008422	11.5
BWI	−0.00813⁺	0.004632	−1.76
DISTDC	−0.18631*	0.016877	−11.04
MHHINC	0.173523*	0.008876	19.55
POPDEN	−0.01219*	0.001726	−7.06
BLPOP	−0.13556*	0.016853	−8.04
CA	−0.27126*	0.010749	−25.24
CH	−0.24072*	0.010283	−23.41
HO	−0.09805*	0.005799	−16.91
LOWRES	−1.10455*	0.320217	−3.45
COMIND	−3.24046*	0.813294	−3.98
MEDHRES	−0.71921*	0.177457	−4.05
CROP	0.011049	0.015922	0.69
FOREST	−1.02805*	0.138842	−7.4
CONSV	0.320958**	0.110929	2.89
PUBLIC	0.082645**	0.031914	2.59
MILIT	0.248354	0.225413	1.1
OTHER	0.297491*	0.012514	23.77
AAPUBLIC	−0.16190*	0.047352	−3.42
AAMILIT	0.553075⁺	0.249905	2.21

*, **, and ⁺ indicate significance at the 0.001, 0.01, and 0.05 levels respectively.

property by $3,307 or 1.87% of the predicted residential value. Conversion of one acre to publicly owned, non-military land use increases the residential value by $994 or 0.57% of the predicted value. Alternatively, a one acre conversion from pastureland to surrounding low density residential land use is found to decrease the value of the mean property by $1,530 or 0.89% of the predicted value and a one acre conversion to commercial/industrial land use decreases the value by $4,450 or 2.56% of the mean residential value. Lastly, a one acre change from pasture to forested land is found to decrease the sales price by $1,424 or 0.82% of the property's value.

As stated earlier, the robustness of these results depends on the validity of the nearest neighbor assumption that was used to determine the minimum allowable distance between any two parcels in the dataset and the success of this technique in controlling for spatial error correlation. To test this, the model is estimated three more times, each time using a randomly drawn sample that uses a successively larger nearest neighbor buffer to exclude neighboring parcels from the dataset. The results from this exercise, using 200, 400, and 600 meters respectively to define the minimum allowable nearest neighbor distance, are reported in Table 3. The results reveal that some, but not all, of

TABLE 3
IV ESTIMATION RESULTS FOR THE 200-, 400-, AND 600-METER SUBSETS

	200-Meter Subset No. Observations: 33,919 R^2: 0.6337 Adj R^2: 0.6334		400-Meter Subset No. Observations: 22,528 R^2: 0.6186 Adj R^2: 0.6181		600-Meter Subset No. Observations: 16,126 R^2: 0.5983 Adj R^2: 0.5976.	
Variable	Estimate	Error	Estimate	Error	Estimate	Error
Intercept	4.725553*	0.36558	4.139167*	0.572787	4.198823*	0.657159
DWGRADE	0.170999*	0.00368	0.173845*	0.004634	0.171793*	0.005728
DWTYPE	0.168842*	0.00749	0.166846*	0.010635	0.157594*	0.013953
BATHS_FU	0.073579*	0.00342	0.07548*	0.004338	0.080146*	0.00539
BATHS_HA	0.061233*	0.0038	0.063075*	0.004905	0.0695*	0.006176
FTPRNT	0.084088*	0.00863	0.08326*	0.010978	0.089392*	0.013597
AREA	0.340489*	0.00789	0.343121*	0.010039	0.333873*	0.012635
LSIZE	0.040997*	0.00248	0.043127*	0.003099	0.044859*	0.003819
AGE	−0.01519*	0.00163	−0.0064**	0.002108	−0.00433+	0.002617
YRSALE	0.026329*	0.0012	0.029726*	0.001561	0.03065*	0.001973
DISTBA	0.109933*	0.00791	0.106342*	0.010078	0.112143*	0.011377
BWI	− 0.01459*	0.00433	−0.02901*	0.005759	−0.03799*	0.007405
DISTDC	−0.18567*	0.01692	−0.16954*	0.031586	−0.18192*	0.042241
MHHINC	0.175677*	0.0081	0.170051*	0.01041	0.173096*	0.01303
POPDEN	−0.01025*	0.00159	−0.00871*	0.002019	−0.00805**	0.002516
BLPOP	−0.10555*	0.01566	−0.09785*	0.020185	−0.08556**	0.025002
CA	−0.26378*	0.00979	−0.23902*	0.012231	−0.23117*	0.015004
CH	−0.23855*	0.00945	−0.23037*	0.012138	−0.23229*	0.015061
HO	−0.09664*	0.00529	−0.10112*	0.006921	−0.10219*	0.008855
LOWRES	−0.14492	0.29882	0.116664	0.283911	−0.02237	0.241808
COMIND	−1.5946+	0.8421	−1.98514	1.404823	−2.41016	1.558216
MEDHRES	−0.1927	0.16642	0.042209	0.14759	−0.00385	0.126872
CROP	0.002865	0.01442	0.005328	0.01762	−0.00713	0.021021
FOREST	−0.83226*	0.13867	−0.70657*	0.211112	−0.77407**	0.267125
CONSV	0.33009**	0.1035	0.255718+	0.125786	0.193644	0.149732
PUBLIC	0.083196**	0.02891	0.05244++	0.035254	0.067251+	0.041522
MILIT	0.277614	0.20044	0.24796	0.229491	0.259036	0.260432
OTHER	0.333589*	0.01146	0.380817*	0.014847	0.406615*	0.018629
AAPUBLIC	−0.21289*	0.04386	−0.23083*	0.05504	−0.2402**	0.066666
AAMILIT	0.488849	0.22561	0.4573++	0.267792	0.313011	0.311205

*, **, +, and ++ indicate significance at the 0.001, 0.005, 0.05, and 0.1 levels respectively.

the land use spillover estimates are robust to this assumption. The coefficients on privately held, preserved open space and public, non-military open space are fairly constant and maintain their significance for the 200 and 400 meter datasets. The privately held, preserved open space estimate loses magnitude and significance with the 600 meter dataset, but the public lands coefficient maintains the same magnitude and is significant at the 0.1 level. In addition, the negative coefficient associated with neighboring forests maintains its significance and decreases in magnitude only slightly and the spillover estimate from commercial/industrial land is negative throughout, although it loses significance with the 400-meter dataset. On the other hand, the estimates of both the low and higher density residential spillovers quickly diminish in magnitude and lose significance. We conclude that the estimates of some of the land use spillovers, most notably those associated with the publicly owned, non-military lands, are robust, others are fairly robust (including the estimates of the preserved lands and commercial/industrial coefficients), and yet others, for example, the estimates of the spillovers from neighboring residential development, are not robust at all.

The more robust estimates from the hedonic pricing model can be used to further explore the implications of the results for open space preservation policies. A question that is particularly relevant for policy concerns the magnitude of the potential benefits from preserving open space. This, of course, depends on the alternative land use—that is, the land use that the parcel would be in if it were not preserved. As reviewed earlier, most of the contingent valuation studies that have elicited open space values specify this alternative land use as being development. However, this confuses the issue by imposing the assumption that the parcel will be developed if it is not preserved. In reality, the alternative land use is usually some type of developable open space (e.g., farmland) that has some positive probability of being converted to a developed use at some point in the future. Therefore, the relevant change that corresponds to the potential benefits associated with farmland preservation is a change in the level of uncertainty associated with the parcel's likelihood of future development. Farmland preservation provides a greater amount of certainty that the land can no longer be developed.

The estimated coefficients associated with the preserved open space variables and the mean values of all explanatory variables are used to calculate the benefits, in terms of the positive spillover effect on neighboring residential property values, associated with preserving a 10-acre plot of farmland.[12] If the land is preserved as an agricultural easement (privately owned conservation land), then the value of a neighboring residential property is estimated to be $4,523 or 2.6% of the predicted mean residential value. Alternatively, if this land is purchased outright by the government and becomes public land, the benefit to the neighboring residential property owner is calculated to be $2,038 or 1.17% of the mean residential value. The difference between the two types of preserved open space, $2,485, is the windfall that the homeowner receives from the surrounding preserved area being privately vs. publicly owned. As discussed earlier, this difference may be attributable to a nuisance spillover associated with public lands. In either case, the homeowner benefits from the preservation of the surrounding farmland. If we consider a case in which more than one residential property is a neighbor to the preserved farmland, then the benefits to preserving the land would clearly increase, depending on the number of neighboring residential properties. In this study area, residential zoning in areas that still have farmland (i.e. exurban or rural areas) typically ranges from 1 to 5-acre minimum lot size. If the 10-acre parcel of farmland were located in the center of this type of residential development, then the predicted benefits if the land was preserved as an agricultural easement range from $10,403 to $52,014 per acre of preserved open space, depending on the density of the neighboring residential development.[13] Alternatively, if the farmland were preserved as publicly owned open space, the predicted benefits range from $4,687 to $23,437 per acre of preserved land, again depending on the density of the surrounding residential development.

These estimates reflect the value associated with the change in the "developable" status of the farmland and are independent of whether the farmland would have been converted to a developed use had it not been preserved. In other words, they correspond to the predicted value of the development rights associated with an open space land parcel. They do not include the additional benefit

[12] This analysis corresponds to the base land use being cropland, pasture, or forest. The estimated coefficients for the preserved open space variables from the models with cropland or forest as the base land use are very similar to those reported in Table 2, in which pasture is the base land use. For the cropland-base model, they are 0.315 and 0.075 for the privately held, preserved open space and public, non-military open space coefficients respectively. Both are significant at the 0.05 level. For the forest-based model, the corresponding estimates are 0.302 and 0.081 and both are significant at the 0.001 level.

[13] Assuming the 10-acre parcel is located within the center of a residential development, 115 residential properties would be within 400 meters of the parcel's centroid if the minimum lot size were 1 acre and 23 properties if the minimum lot size were 5 acres.

from avoiding the negative spillovers that otherwise would be generated by the parcel being in a developed use.

CONCLUSIONS

Results from this analysis show that surrounding open space significantly influences the residential sales price of houses and that different types of open space have differing effects. Using techniques aimed at addressing the identification problems that arise from endogenous variables and unobserved spatial correlation, we find that the spillover effects from preserved open space are significantly greater than those associated with developable farmland and forest. In addition, the findings show that the spillovers from pasture vs. cropland are not significantly different from one another, but that pastureland does generate a significantly greater spillover effect on residential property values than does the spillover effect of neighboring forests. Other results, such as the negative spillovers associated with neighboring residential, are not found to be robust to different assumptions regarding the extent of the spatial error correlation.

The validity of these results rests on the success of the econometric techniques used to solve the identification problems that arise from unobserved spatial correlation and endogenous variables. An instrumental variables technique was used to address the problems of bias associated with the land use spillover estimates. In addition, because this approach fails to correct for the remaining spatial error correlation, a spatial sampling technique was used to eliminate nearest neighbors from the dataset. To the extent that the instruments are poor (e.g., they are not strongly correlated with the land use measures or they are correlated with the error) or to the extent that the spatial error correlation is not adequately controlled using the sample that excludes nearest neighbors, the results will be incorrect. Experimentation with the definition of the minimum nearest neighbor distance used to select the sample revealed that some of the results are not robust to these assumptions. The lack of robustness that was evident in some of these estimates illustrates the challenges involved in solving the identification problems that arise due to endogeneity and spatial error correlation, an issue that has not been adequately addressed in the literature.

Other limitations of the analysis include the fact that the results are not appropriate for describing the welfare effects of non-marginal changes in the regional housing market and the possibility that a sample selection problem may be present. This latter problem arises because we only have data on houses that were actually transacted and not on those that were not transacted. To the extent that there may be systematic differences between houses that were transacted vs. those that were not, this would lead to biased results. Finally, the estimated marginal values associated with the different types of open space investigated in this paper do not capture all of the potential values associated with open space. For example, they do not include the value that non-property owners may ascribe to open space, including the values held by visitors to public parks. In addition, they do not include any nonuse values associated with open space, e.g., the role that open space plays in protecting groundwater, wildlife habitat, and natural places.

Despite these limitations, the findings do provide some useful insights regarding the demand for open space preservation. First, the results shed some light on the specific attributes of open space that are most valued and the extent to which open space may be most valued for simply not being development. The evidence suggests that the public's demand for open space preservation is motivated more by the fact that open space implies no development rather than being driven by particular features of open space landscapes. Specifically, we find that significant additional benefits are estimated to accrue to neighboring residential properties given a marginal change in the landscape from any of the developable open spaces considered here (cropland, pasture, and forest) to either of the two preserved open space uses, privately owned land in conservation or publicly owned, non-military land. In contrast, a marginal conversion from pasture to cropland is not found to generate significantly different spillover effects. While these results support the hypothesis that it is the absence of development that primarily drives the demand for open space, there is limited evidence that supports the hypothesis that particular features of open space landscapes also matter—namely, the negative and significant effect of a marginal change in the landscape from pasture to forests is found to be robust.

Secondly, the results provide a partial estimate of the total marginal benefits from open space preservation and therefore offer some guidance for the design of open space preservation policies. The benefits of preserving any particular piece of open space are a function of the number of residents within the neighboring area, their preferences, and the relative scarcity of open space in the region. Given the conditions within the Maryland study area, which are characterized by relatively high rates of population growth and land conversion, we find that the marginal benefits to one household of preserving neighboring open space range from $994 to $3,307 per acre of farmland that is preserved, depending on whether it is publicly or privately owned. Aggregated across the total number of households that are within close enough proximity to benefit from the positive spillover and across the total number of acres preserved, these numbers provide a lower bound of the benefits that accrue to a community from open space preservation. As such, they offer guidance to local decision makers in regions with similar growth conditions who are seeking to quantify the costs and benefits associated with open space preservation programs.

References

Anselin, L. 1988. *Spatial Econometrics: Methods and Models.* The Netherlands: Kluwer Academic Publishers.

Beasley, S. D., W. G. Workman, and N. A. Williams. 1986. ''Estimating Amenity Values of Urban Fringe Farmland: A Contingent Valuation Approach.'' *Growth and Change* 17:70–78.

Bergstrom, J., B. Dillman, and J. Stoll. 1985. ''Public Environmental Amenity Benefits of Private Land: The Case of Prime Agricultural Land.'' *Southern Journal of Agricultural Economics* 17 (1): 139–49.

Bower, J., and D. Didychuk. 1994. ''Estimation of Market Benefits of Agricultural Land Retention in Eastern Canada.'' *Agricultural and Resource Economics Review* 23 (2): 218–55.

Brock, W., and S. Durlauf. 2001. ''Interactions-Based Models.'' In *The Handbook of Econometrics 5,* ed. James J. Heckman and Edward E. Leamer. Amsterdam: Elsivier Science, North Holland.

Cheshire, P., and S. Sheppard. 1995. ''On the Price of Land and the Value of Amenities.'' *Economica* 62 (May): 247–67.

Cropper, M. L., L .B. Deck, and K. E. McConnell. 1988. ''On the Choice of Functional Form for Hedonic Price Functions.'' *Review of Economics and Statistics* 70 (4): 668–75.

Freeman, M. 1993. *The Measurement of Environmental and Resource Values.* Washington D.C.: Resources for the Future.

Garrod, G., and K.Willis. 1992. ''The Environmental Economic Impact of Woodland: A Two-Stage Hedonic Price Model of the Amenity Value of Forestry in Britain.'' *Applied Economics* 24 (7): 715–28.

Geoghegan, J. 2002. ''The Value of Open Spaces in Residential Land Use.'' *Land Use Policy* (1): 91–98.

Geoghegan, J., L. A. Wainger, and N. E. Bockstael. 1997. ''Spatial Landscape Indices in a Hedonic Framework: An Ecological Economics Analysis Using GIS.'' *Ecological Economics* 23 (3): 251–64.

Greene, W. H. 2000. *Econometric Analysis*, 4th ed. Upper Saddle River, N.J.: Prentice Hall.

Haining, R. 1993. *Spatial Data Analysis in the Social and Environmental Sciences.* New York: Cambridge University Press.

Halstead, J. 1984. ''Measuring the Nonmarket Value of Massachusetts Agricultural Land: A Case Study.'' *Journal of Northeastern Agricultural Economic Council* 13 (1): 226–47.

Heimlich, R., and W. Anderson. 2001. ''Development at the Urban Fringe and Beyond: Impacts on Agricultural and Rural Land.'' Economic Research Service, U.S. Department of Agriculture. AER No. 803.

Irwin, E .G., and N. E. Bockstael. 2001a. ''The Problem of Identifying Land Use Spillovers:

Measuring the Effects of Open Space on Residential Property Values." *American Journal of Agricultural Economics* 83 (3): 698–704.

———. 2001b. "Interacting Agents, Spatial Externalities, and the Endogenous Evolution of Residential Land Use Pattern." *Journal of Economic Geography* 2 (1): 31–54.

Krieger, D. 1999. "Saving Open Spaces: Public Support for Farmland Protection." Working paper. Center for Agriculture in the Environment, American Farmland Trust.

Manski, C. 1993. "Identification of Endogenous Social Effects: The Reflection Problem." *Review of Economic Studies* 60 (3): 531–42.

Nelson, G., and D. Hellerstein. 1997, "Do Roads Cause Deforestation? Using Satellite Images and Econometrics of Land Use." *American Journal of Agricultural Economics* 79 (1): 80–88.

Tyrävinen, L., and A. Miettinen. 2000. "Property Prices and Urban Forest Amenities." *Journal of Environmental Economics and Management* 39 (2): 205–23.