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The effects of potential land development on agricultural land prices

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Abstract

We conduct a national-scale analysis of the determinants of agricultural land values. The theoretical basis for the study is a spatial city model with stochastic returns to future land development. The empirical model of agricultural land prices is estimated with a cross-section on approximately three thousand counties in the contiguous US. The results provide evidence that option values associated with irreversible and uncertain land development are capitalized into current farmland values. For each county, we decompose the current agricultural land value into components measuring rents from agricultural production and rents from future land development.

JEL classification: Q11; Q15; R14

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0. Introduction

Land prices reflect not only the uses of land, but the potential uses as well. In a competitive market, the price of land will equal the discounted sum of expected net returns obtained by allocating the land to its most profitable use. That use surely may change over time. If, for example, agricultural production is currently the most profitable use, but development for some other purpose is expected to yield even greater net returns in the future, then the current land price should reflect the returns to both uses in a simple additive form: the sum of the discounted stream of rents from agriculture up until the time of conversion plus the discounted stream of expected rents from development from that time onward.

For many years, economists have analyzed the structure of agricultural land prices in an effort to understand potential threats to agriculture posed by land development and to identify policies to prevent or discourage what may be considered to be socially undesirable land-use changes. In the United States, the loss of agricultural land to urbanization has been an enduring policy issue because of concerns that a reduced domestic capacity to produce food could threaten national security and because of losses of open space and other environmental amenities in rapidly urbanizing areas. In 1973, President Richard Nixon proclaimed farmland protection to be the nation's most pressing environmental issue. In 1979, the US Secretary of Agriculture Robert Bergland warned that "Continued destruction of cropland is wanton squandering of an irreplaceable resource that invites tragedy not only nationally, but on a global scale." Recently, the 1996 Federal Agricultural Improvement and Reform Act expanded the federal role in agricultural land preservation by funding the purchase of farmland conservation easements. In the last decade, there has been rapid growth in the number of private land trusts in the US, many of which are devoted to preserving agricultural land through the purchase of development rights.

Previous studies have examined the effects of population, income, and other determinants of development rents on farmland prices, but have been unable to separate the contributions to market value of rents due to agricultural use and rents due to potential development.[4] Decomposing farmland prices into their additive components can be of considerable value to understanding potential development

[4] Earlier analyzes of farmland prices that include proxy variables for future development rents are Hushak and Sadr [15], Chicoine [8], Shonkwiler and Reynolds [30], Palmquist and Danielson [23], Elad et al. [12], Mendelsohn et al. [21], Vitaliano and Hill [34], Shi et al. [29], and Plantinga and Miller [24]. Hardie et al. [14] estimate a model in which average farmland and housing prices are simultaneously determined, and include income, population, and accessibility variables as exogenous determinants of housing rents. There are also a large number of related analyzes of the determinants of developed land prices (for example, Coulson and Engle [11], Peiser [25], Kowalski

paths, because high current land prices may reflect profitable current use, potential for a more profitable use in the future, or some combination of both. In areas where high current prices are found to be largely a result of capitalized rents from future land development, market intervention may be warranted to prevent losses of agricultural land and associated public benefits. A major obstacle to such price component identification has been the obvious unobservability of the date of future development. Complicating matters further is the likely presence of option values associated with the land development decision. Because of uncertainty about future returns to development and the prohibitively high cost of reversing farmland conversion, there may be considerable value to preserving the option to develop land. In general, option values affect both the timing of land conversion and the current price of farmland.[5]

In this paper, we seek to better understand the dynamic structure of land prices by estimating a model of farmland values that explicitly accounts for uncertainty over future development rents and allows decomposition of the current value into agriculture and development components.[6] In the theoretical model of a land market presented in Section 1, future development rents are assumed to evolve according to a specified stochastic process. By imposing this structure on development rents, we can solve for the expected conversion time. Incorporating this result, equilibrium prices for agricultural land become a function of the expected growth rate and variance of development rents, for which suitable proxy variables can be obtained. The econometric analysis described in Section 2 draws upon county-level data for the forty-eight contiguous United States. Using the theoretical model, we derive an expression for the average price of agricultural land in a county in terms of average agricultural returns, average prices of recently developed land, and other observable variables. The agriculture and development components of the average farmland price are identified in this expression and, thus, can be recovered from the estimated econometric model. We present the empirical results in Section 3, including estimates of the development component's share of agricultural land prices for all counties in the contiguous US. Finally, in Section 4, we conclude with a discussion of policy implications, with particular emphasis on how the results can inform the development of farmland preservation policies.

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and Paraskevopoulos [18], Rosenthal and Helsley [27], Colwell and Munneke [10], McDonald and McMillen [20]).

[5] Theoretical and empirical studies that consider the effect of option values on land development and land prices include Titman [32], Quigg [26], Capozza and Sick [7], Tegene et al. [31], Capozza and Li [6], and Turvey [33].

[6] In addition, our study advances the methodology on analyzing farmland prices by providing a stronger theoretical motivation for the variables in our empirical model, explicitly accounting for the aggregate structure of our data in the model specification, and using a more reliable measure of development rents.

1. The theoretical model

In a competitive market, where risk-neutral landowners seek to maximize the economic returns to their land, the market price of an agricultural parcel at time t that will be developed at t^* will be equivalent to the present discounted value of the stream of expected net agricultural returns from time t to t^* (the agriculture component) plus the present discounted value of the stream of expected net returns from the developed parcel subsequent to time t^* (the development component)

$$P_t^{\mathrm{A}}(t^*, \mathbf{z}) = E_t\left\{\int_t^{t^*} \pi^{\mathrm{A}}(s, \mathbf{z})\,\mathrm{e}^{-r(s-t)}\,\mathrm{d}s + \int_{t^*}^{\infty} \pi^{\mathrm{D}}(s, \mathbf{z})\,\mathrm{e}^{-r(s-t)}\,\mathrm{d}s - C\,\mathrm{e}^{-r(t^*-t)}\right\}, \tag{1}$$

where $\pi^{\mathrm{A}}(s, \mathbf{z})$ is the net return to agriculture at time s and location $\mathbf{z}$, where $\mathbf{z}$ is a two-dimensional vector of spatial coordinates, $\pi^{\mathrm{D}}(s, \mathbf{z})$ is the net return to development at time s and location $\mathbf{z}$, r is the discount rate (presumably a function of the anticipated rate of return on alternative investments), and C is the cost of developing agricultural land. The expression in (1) reflects the assumption that land development is irreversible.

We impose additional structure on (1) in order to produce an estimable model of agricultural land prices. Except for minor alterations, the model presented in the rest of this section is identical to Capozza and Helsley's [5] urban spatial model with stochastic development rents. First, we specify agricultural rents as $\pi^{\mathrm{A}}(s, \mathbf{z}) = \pi^{\mathrm{A}}$ for all s, where π^{A} is the average agricultural rent in the vicinity of $\mathbf{z}$.[7] Net returns to agricultural are assumed to be constant over time. Alternative specifications allowing for linear rates of change or exponential growth in returns lead to an analytically intractible model of land prices. Fortunately, for the empirical application presented in the next section, there is some justification for the assumption of constant agricultural returns, a point we elaborate on below.

Second, we specify the rents from land development as $\pi^{\mathrm{D}}(s, \mathbf{z}) = m_1(s) + m_2(\mathbf{z})$. A common feature of urban spatial models is a bid rent for developed land that declines in distance from a center of economic activity such as a central business district (CBD).[8] Hence, we specify the spatial component of development rents as $m_2(\mathbf{z}) = -\gamma z$, where γ is a positive parameter and z is the distance from the CBD. The temporal component of development rents is specified as $m_1(s) \equiv gs + \sigma B(s)$, where $B(s)$ is a standard Brownian motion

[7] The $\mathbf{z}$ argument is dropped to simplify the notation. It should be understood that average agricultural rents are specific to a geographical area (counties, in the econometric analysis).

[8] See, for example, Mills [22] and Capozza and Helsley [4].

process with zero drift and variance 1, and g and σ are positive parameters (that is, $m_1(s)$ follows a Brownian motion process with drift g and variance σ^2). Additive separability of the spatial and temporal components implies that the rate of change in development rents is independent of location. Capozza and Helsley [5] show that this specification follows from a simple household equilibrium model in which households consume a fixed amount of land and income is subject to exogenous shocks.

The basic statistical properties of $m_1(s)$ carry over to the development rent function $\pi^{\mathrm{D}}(s, z)$. In particular, it follows that:

$$\pi^{\mathrm{D}}(t+s', z) \xrightarrow{d} \pi^{\mathrm{D}}(t, z) + gs' + \sigma B(s'), \tag{2}$$

which indicates that the distribution of development rents s' periods in the future is equivalent to that of the current (time t) development rents plus the drift and random components evaluated at s'. As such, the expected value of development rents at any time in the future is conditional on current development rents. The result in (2) can be used to solve the second integral in (1). Specifically, we can write the expectation of a discounted stream of development rents beginning in period $t^* > t$ as

$$\begin{aligned} &E\left\{\int_{t^*}^{\infty}\left[\pi^{\mathrm{D}}(t^*, z) + g\cdot(s-t^*) + \sigma B(s-t^*)\right]\mathrm{e}^{-r(s-t)}\,\mathrm{d}s \;\Big|\; \pi^{\mathrm{D}}(t, z)\right\} \\ &\quad = E\left\{\left[\frac{\pi^{\mathrm{D}}(t^*, z)}{r} + \frac{g}{r^2}\right]\mathrm{e}^{-r(t^*-t)} \;\Big|\; \pi^{\mathrm{D}}(t, z)\right\}, \end{aligned} \tag{3}$$

where the derivation of the right-hand side term makes use of integration by parts and $E_t[\sigma B(t+s)] = 0$.

Substituting (3) into (1) and incorporating the specification of agricultural rents from above, the price of agricultural land at time t is written

$$\begin{aligned} P_t^{\mathrm{A}}(t^*, z) = {} & \frac{\pi^{\mathrm{A}}}{r} E\left\{1 - \mathrm{e}^{-r(t^*-t)} \;\big|\; \pi^{\mathrm{D}}(t, z)\right\} \\ & + E\left\{\left[\frac{\pi^{\mathrm{D}}(t^*, z)}{r} + \frac{g}{r^2} - C\right] e^{-r(t^*-t)} \;\Big|\; \pi^{\mathrm{D}}(t, z)\right\}, \end{aligned} \tag{4}$$

where the first and second terms are respectively the agriculture and development components of the current land price. A risk-neutral landowner seeking to maximize the economic returns to his land will choose t^* to maximize $P_t^{\mathrm{A}}(t^*, z)$. This can be solved as an optimal stopping problem in which the landowner develops the parcel at the first time development rents reach a reservation value $R^* = \pi^{\mathrm{D}}(t^*, z)$ that compensates him for agricultural returns, the opportunity cost of land conversion, and an option value related to the foregone opportunity to further delay the irreversible land development decision. The random component of price in this problem is the stopping time, $t^* - t$. From Karlin and Taylor [16, pp. 361–362] the expected value of the Laplace transform of the stopping time conditional on the initial value of development rents, $\pi^{\mathrm{D}}(t,z)$, and the reservation value, R^*, is given by

$$E\bigl\{\mathrm{e}^{-r(t^*-t)} \mid \pi^{\mathrm{D}}(t,z), R^*\bigr\} = \mathrm{e}^{-\alpha[R^*-\pi^{\mathrm{D}}(t,z)]}, \tag{5}$$

where $\alpha = [(g^2+2\sigma^2 r)^{1/2} - g]/\sigma^2$. Substituting (5) into (4) gives the price of agricultural land at location z

$$P_t^{\mathrm{A}}(z) = \frac{\pi^{\mathrm{A}}}{r}\bigl(1-\mathrm{e}^{-\alpha[R^*-\pi^{\mathrm{D}}(t,z)]}\bigr) + \left[\frac{R^*}{r} + \frac{g}{r^2} - C\right]\mathrm{e}^{-\alpha[R^*-\pi^{\mathrm{D}}(t,z)]}. \tag{6}$$

Landowners will choose the reservation value R^* to maximize the value of the land. Differentiation of (6) with respect to R^* yields the optimal reservation value $R^* = \pi^{\mathrm{A}} + rC + (r-\alpha g)/\alpha r$, where $r - \alpha g \geqslant 0$.

The declining rent gradient for developed land implies that land close to the CBD will be developed first. In time t, all parcels at distance $z^*(t)$ will be developed where $z^*(t) = [m_1(t) - R^*]/\gamma$ from the definition of $\pi^{\mathrm{D}}(s,z)$. Using the definitions of $z^*(t)$ and $\pi^{\mathrm{D}}(t,z)$, $R^* - \pi^{\mathrm{D}}(t,z)$ can be expressed as $\gamma[z - z^*(t)]$ and the price of agricultural land at location $z > z^*(t)$ can be rewritten as

$$P_t^{\mathrm{A}}(z) = \frac{\pi^{\mathrm{A}}}{r}\bigl(1-\mathrm{e}^{-\alpha\gamma[z-z^*(t)]}\bigr) + \left[\frac{\pi^{\mathrm{A}}}{r} + \frac{1}{\alpha r}\right]\mathrm{e}^{-\alpha\gamma[z-z^*(t)]}, \tag{7}$$

where $(1/\alpha r)\,\mathrm{e}^{-\alpha\gamma[z-z^*(t)]}$ is the option value associated with delaying land conversion. The price of a parcel of land developed at time t is equal to the expected present discounted value of the stream of development rents from time t onward, and can be written

$$P_t^{\mathrm{D}}\bigl(z^*(t)\bigr) = \frac{\pi^{\mathrm{A}}}{r} + \frac{1}{\alpha r} + C. \tag{8}$$

Eq. (8) is derived by substituting the optimal reservation value R^* for $\pi^{\mathrm{D}}(t^*,z)$ and t^* for t in the right-hand side of (3). Importantly for the empirical analysis presented below, Eqs. (7) and (8) can be combined to yield an expression for the current price of agricultural land in terms of the price of a parcel developed in the current period

$$P_t^{\mathrm{A}}(z) = \frac{\pi^{\mathrm{A}}}{r}\bigl(1-\mathrm{e}^{-\alpha\gamma[z-z^*(t)]}\bigr) + \bigl[P_t^{\mathrm{D}}\bigl(z^*(t)\bigr) - C\bigr]\mathrm{e}^{-\alpha\gamma[z-z^*(t)]}, \tag{9}$$

where, as above, the first term is the agriculture component and the second is the development component. The option value is now subsumed in the development component.

2. Econometric estimation

The price expression in (9) serves as the theoretical basis for an econometric analysis conducted with data on all counties in the contiguous forty-eight United States. We have, for each county, an estimate of the average per-acre price of agricultural land in 1997. In terms of (9), these data represent an average of $P_t^{\mathrm{A}}(z)$ over the undeveloped parcels in the county, comprising the area between the city boundary and the county boundary. Formally, if $\bar{z}_i$ is the distance from the CBD to the boundary of county i, the average price of agricultural land is given by

$$\overline{P}_{it}^{\mathrm{A}} = \int_{z_{it}^*}^{\bar{z}_i} \left(\pi_i^{\mathrm{A}} \cdot \frac{1 - \mathrm{e}^{-\alpha_i \gamma_i [z - z_{it}^*]}}{r(\bar{z}_i - z_{it}^*)} + \left[P_{it}^{\mathrm{D}}(z_{it}^*) - C \right] \cdot \frac{\mathrm{e}^{-\alpha_i \gamma_i [z - z_{it}^*]}}{\bar{z}_i - z_{it}^*} \right) \mathrm{d}z, \quad (10)$$

where the parameters α_i and γ_i are assumed to vary across counties.[9] Eq. (10) shows that the current average price of agricultural land can be expressed in terms of the net return to agriculture (π_i^{A}), the current price of recently developed land ($P_{it}^{\mathrm{D}}(z_{it}^*)$), the rate of change in development rents (g_i), the variance of shocks to development rents (σ_i^2), the rate of change in development rents as distance to the CBD increases (γ_i), and the remaining amount of agricultural land in the county ($\bar{z}_i - z_{it}^*$).[10] Conversion costs (C) and the interest rate (r) are assumed to be constant across counties.

Comparative statics results indicate that $\overline{P}_{it}^{\mathrm{A}}$ is increasing in π_i^{A}, $P_{it}^{\mathrm{D}}(z_{it}^*)$, g_i, and σ_i^2, and decreasing in γ_i and $(\bar{z}_i - z_{it}^*)$.[11] All else equal, increases in agricultural rents and the current price of developed land raise the agriculture and development components of the land price, respectively. A greater rate of change in development rents implies larger future returns to developed land, which are capitalized into the current land price. A higher variance in development rent shocks increases the option value associated with delaying the irreversible land conversion decision. A larger value of γ_i implies that development rents fall off more quickly as distance to the CBD increases, reducing the value of capitalized development rents. Finally, more agricultural land dilutes the effects of future development rents on the average price of agricultural land.

[9] While (10) relies on a highly stylized model of urban and rural land use in a county—in particular, the county is assumed to be circular with the CBD located at its center—solution of the integral indicates that the average agricultural land price depends on $(\bar{z}_i - z_{it}^*)$, which, more generally, indicates how much agricultural land remains in the county.

[10] Note that g_i and σ_i^2 are subsumed in α_i in Eq. (10).

[11] These results are most easily established from (7). For the last result, note that increasing $(\bar{z}_i - z_{it}^*)$ is equivalent to adding agricultural land beyond the county boundary. This must lower the average price of agricultural land because, from (7), $\partial P_t^{\mathrm{A}}(z)/\partial z < 0$.

The model is estimated with a cross-section on $N = 2955$ counties in 1997.[12] Solving the integral in (10) and suppressing time subscripts and arguments, the empirical model is written

$$\overline{P}_i^{\mathrm{A}} = \beta_{0i} + \beta_{1i}\pi_i^{\mathrm{A}} + \beta_{2i}P_i^{\mathrm{D}} + u_i, \quad i = 1, \ldots, N, \tag{11}$$

where

$$\begin{aligned} \beta_{0i} &= -(\bar{z}_i - z_i^*)^{-1}C\big(1 - \mathrm{e}^{-\alpha_i\gamma_i(\bar{z}_i - z_i^*)}\big), \\ \beta_{1i} &= \big[(\bar{z}_i - z_i^*)r\big]^{-1}\big(\bar{z}_i - z_i^* - (\alpha_i\gamma_i)^{-1}\big)\big(1 - \mathrm{e}^{-\alpha_i\gamma_i(\bar{z}_i - z_i^*)}\big), \\ \beta_{2i} &= (\bar{z}_i - z_i^*)\big(1 - \mathrm{e}^{-\alpha_i\gamma_i(\bar{z}_i - z_i^*)}\big), \end{aligned} \tag{12}$$

and u_i is a random disturbance whose statistical properties are discussed below. Written this way, the land price model has an intuitive interpretation. The price of agricultural land is seen to depend on a weighted sum of the net returns to agriculture and the current price of developed land. From the underlying theory, the weights depend on the rate of change in and variance of development rents, accessibility, and the stock of agricultural land.

Because, in general, the weights on π_i^{A} and P_i^{D} will vary across counties, the number of parameters in (11) exceeds the number of observations and (11) does not represent a feasible estimation problem. One solution is to restrict the βs to be equal across counties. The estimation results for this version of the model are presented below for the purpose of comparison. However, we focus our attention on a less restrictive model in which the parameters β_{i0}, β_{i1}, β_{i2} as specified as functions of additional variables and parameters that are constant across the set of counties. Since we do not know the exact relationship between the βs and the independent variables, we approximate the relationship with the quadratic function[13]

$$\begin{aligned} \beta_{ji} = {} & c_{j0} + c_{j1}\mathit{cpopd}_i + c_{j2}(\mathit{cpopd}_i)^2 + c_{j3}\mathit{vpopd}_i + c_{j4}(\mathit{vpopd}_i)^2 \\ & + c_{j5}\mathit{roads}_i + c_{j6}(\mathit{roads}_i)^2 + c_{j7}\mathit{farms}_i + c_{j8}(\mathit{farms}_i)^2, \end{aligned} \tag{13}$$

for $j = 0, 1, 2$ and where cpopd_i, vpopd_i, roads_i, and farms_i are proxies, discussed below, for g_i, σ_i^2, γ_i, and $(\bar{z}_i - z_i^*)$, respectively. Substitution of (13) into (11) yields a feasible estimation problem and allows the βs to take different values in each county.

Details on the data used to estimate the model are found in Appendix A. All variables are for county i and the year 1997 unless otherwise indicated. $\overline{P}_i^{\mathrm{A}}$ is

[12] One hundred fifty-six counties are omitted because of missing data or the absence of agricultural land.

[13] This parsimonious specification was selected over a more general polynomial function (for example, Plantinga and Miller [24]) because of collinearity between interaction and higher-order terms.

the average per-acre estimated value of farmland. π_i^{A} is the per-acre average net return from agricultural land, including federal farm subsidies. A principal goal of federal farm programs is to stabilize farm incomes through price supports. For example, for one program affecting major farm commodities such as corn, wheat, and soybeans, farmers can receive deficiency payments equal to the difference between a target price and the market price. Since the early 1980s, target prices for most program crops have typically exceeded market prices, indicating that revenues, at least, have been relatively constant. The influence of farm programs on crop prices lends support to our assumption—here and in the theoretical model—of constant net returns to agriculture.[14]

P_i^{D} is a county-level estimate of the average per-acre price of recently developed land. This variable measures the average value of a developed parcel less the value of structures, and thus corresponds to the present discounted value of the stream of rents from improved bare land. Improvements may include sewer lines, driveways, and landscaping. The costs of these improvements are captured in the conversion cost term (C).

Historical population statistics are used to develop proxy measures for the growth rate and variance of changes in future development rents (respectively, g_i and σ_i^2). For this empirical application, we need to account for potential differences across counties in future rents to developed land, and many of the factors that determine these differences are subsumed in expectations of population growth. For example, a demand shock that increases labor demand in one region will increase migration to the region (provided the costs of migration are not too great), and the influx of migrants will bid up rents for developed land. Participants in the land market are assumed to form expectations of future population changes based on recent past changes. The average annual change in total county population density between 1990 and 1997 (denoted $cpopd_i$) is used as a proxy measure for g_i, and the variance of annual changes in population density over the same period (denoted $vpopd_i$) proxies for σ_i^2.[15]

In spatial city models, development rents typically fall with distance to the CBD in order to compensate residents for higher commuting costs. Thus, one reason why γ_i, the "spatial rate of change" in development rents, might vary across counties is differences in travel costs. We use highway road density in a county ($roads_i$) as a proxy measure for γ_i. Higher road density is expected to

[14] One could argue that, in 1997, farmers would have expected future net returns to decline due to the phase-out of price supports specified under the 1996 farm legislation. However, with the benefit of hindsight, ones sees that farm subsidies remained at historical levels through the late 1990s and have been recently restored as part of the 2002 Farm Bill.

[15] Capozza and Helsley's [5] analysis, upon which our econometric model is based, is of an open city model with costless migration. In such models, population is determined endogenously. In our empirical analysis $cpopd_i$ and $vpopd_i$ are included as exogenous determinants of future development rents. These variables are proxy measures for *ex post* changes in development rents, which are assumed to form the basis for expectations of future changes in these rents.

reduce γ_i, and based on the comparative statics results discussed above, increase the land price. The remaining area of agricultural land $(\bar{z}_i - z_i^*)$ is measured as total farmland acres ($farms_i$) divided by the county land area.

The remaining estimation issue is the statistical properties of the error term in (11). Given that our data are cross-sectional and spatially-referenced, we allow for a heteroskedastic and spatially-correlated[16] error structure

$$\mathbf{u} = \rho \mathbf{W}\mathbf{u} + \mathbf{e}, \qquad e_i \sim \left(0, v_i^2\right), \tag{14}$$

where $\mathbf{u}$ and $\mathbf{e}$ are vectors of random variables, ρ is a scalar, $\mathbf{W}$ is an $N \times N$ weight matrix indicating the spatial structure of the data, and v_i^2 is the variance of the ith element (e_i) of $\mathbf{e}$. Standard tests (for example, White's [35] test) reject the null hypothesis of homoskedasticity. To adjust the residuals for heteroskedasticity, we assume that the error variance is an increasing function of the county land value.[17] Since we do not know the precise relationship between land values and the corresponding error variance, we begin by dividing the data into deciles according to the magnitude of the reported land value. For each group (approximately 300 observations), we compute an estimate of the error variance. The estimated error variances are similar in magnitude for the lower six deciles, but then increase considerably with higher land values. The variance estimates are used to weight the data and the model is re-estimated using the feasible GLS estimator.[18]

We test for spatial autocorrelation using Moran's I statistic $I = N(\hat{\mathbf{e}}'\mathbf{W}\hat{\mathbf{e}})/S(\hat{\mathbf{e}}'\hat{\mathbf{e}})$, where $\hat{\mathbf{e}}$ is the N-vector of estimated residuals, and S is a standartization factor equal to the sum of the elements of $\mathbf{W}$.[19] Computation of Moran's I statistic requires knowledge of $\mathbf{W}$. In particular, we must specify which non-diagonal elements of the variance–covariance matrix are non-zero and the weights (if any) on each of these elements. Common practice is to assume non-zero covariances for counties that share a common border. In this case, each element of $\mathbf{W}(w_{ij})$

[16] Since we model only within-county effects of the independent variables, a potential source of spatial autocorrelation is cross-county effects of these variables on land values.

[17] In counties with large land values, a greater share of the value is likely to be determined by future rents from development, which are unobserved and speculative. In contrast, in counties with small land values, most of the land value is derived from relatively certain agricultural returns. In addition, the magnitude of data reporting and compilation errors is larger in counties with high land values if these errors are proportional to the quantities being measured.

[18] Even after the heteroskedasticity correction, large positive residual estimates are found for counties in Connecticut, Massachusetts, and New Jersey. Since agricultural land values are very high in these cases, we suspect that our model is failing to capture some of the factors determining future development rents in these counties. Accordingly, separate dummy variables were included in the model for each of these states.

[19] Moran's I is a spatial analogue to Pearson's correlation coefficient. It takes values between -1 (strong negative autocorrelation) and 1 (strong positive autocorrelation) in most applications (Bailey and Gatrell [3]) and under the null hypothesis of no spatial autocorrelation has an expected value of $-1/(N-1)$, which converges to zero as N increases. See Anselin [2] for a detailed discussion of Moran's I.

takes a value 1 if county i is adjacent to county j and is 0 otherwise. The computed value of Moran's I is 0.54, indicating fairly strong spatial autocorrelation.[20] Assuming an approximate standard normal distribution for Moran's I statistic, the corresponding z statistic is approximately 51, and so the null hypothesis of no spatial autocorrelation is rejected at any reasonable confidence level.

To adjust the residuals for spatial autocorrelation, we must estimate the spatial autoregressive parameter ρ. We use the generalized moments estimator developed by Kelejian and Prucha [17]. This approach is particularly suited for this application, as other available estimators may not be computationally feasible in cases with large numbers of observations. Applying Eq. (7) in Kelejian and Prucha, we form an estimate of ρ and transform the data using the matrix $\widehat{\mathbf{P}} = \mathbf{I}_N - \hat{\rho}\mathbf{W}$, where $\mathbf{I}_N$ is an N-dimensional identity matrix. The corresponding feasible GLS estimates are then computed.

3. Results

The model of agricultural land values appears to have a good fit,[21] and most of the coefficient estimates, including many second-order terms, are significantly different from zero at the 5% level (Table 1). Since the signs and magnitudes of individual coefficients do not have clear interpretations, we compute the partial effects of π^{A}, P^{D}, *cpopd*, *vpopd*, *roads*, and *farms* on $\overline{P}^{\mathrm{A}}$ and evaluate the resulting expressions at the estimated coefficient values and means of the other independent variables (Table 2). Standard errors are computed using the delta method. All of the partial effects are significantly different from zero at the 5% level and, except in one case, have the signs indicated by the comparative statics results discussed above.

In the average county, a \$1 increase in the annual per-acre return to agriculture (π^{A}) increases the value of agricultural land by \$5.00.[22] A \$1 increase in the current price of developed land (P^{D}) decreases the agricultural land value by

[20] The elements of **W** were generated with ArcInfo, a spatial data analysis program, and I was computed with an algorithm programmed by the authors.

[21] The adjusted R^2 measure has a limited interpretation in the GLS context; it indicates that the transformed independent variables explain 67% of the variation in the transformed dependent variable.

[22] It is tempting to use this result to compute the implicit time of development. The present value of a series of \$1 payments terminating in year n is given by $[(1+r)^n - 1]/r(1+r)^n$, implying in our case that $n = 6$ when $r = 5\%$. However, caution must be used in interpreting n, as this is the development for the average parcel as opposed to the average development time across parcels. If $f(t^*)$ is the density function of optimal development times for all parcels in the US, then, in continuous time, $n = -\ln[\int_0^\infty e^{-rt^*} f(t^*)\,\mathrm{d}t^*]/r$. It can be shown that n is always less than the average development time given by $\bar{n} = \int_0^\infty t^* f(t^*)\,\mathrm{d}t^*$. Indeed, the divergence between n and $\bar{n}$ can be considerable. Suppose that $f^*(t)$ is a discrete uniform distribution on $[1, 200]$ and the interest rate is 5%. Then, $\bar{n} = 100.5$ years and $n = 46.6$ years.

Table 1
Feasible generalized least squares estimates for the land value model

Variable	Coefficient estimate	Standard error
Intercept	867.93[a]	35.57
cpopd	78.84[a]	7.55
cpopd ^{2}	−1.02[a]	0.14
vpopd	−0.03	0.07
vpopd^2	1.8E–06[a]	3.79E–07
roads	1001.32[a]	208.23
roads^2	−696.76[a]	299.03
farms	320.86[a]	152.2
farms^2	−862.44[a]	147.75
π^{A}	1.86[a]	0.41
$\pi^{\mathrm{A}} \cdot cpopd$	-0.01	0.05
$\pi^{\mathrm{A}} \cdot cpopd^2$	2.25E–03	1.42E–03
$\pi^{\mathrm{A}} \cdot vpopd$	−2.93E–03[a]	6.52E–04
$\pi^{\mathrm{A}} \cdot vpopd^2$	5.12E–09[a]	1.18E–09
$\pi^{\mathrm{A}} \cdot roads$	−4.12[a]	1.79
$\pi^{\mathrm{A}} \cdot roads^2$	6.40[a]	2.43
$\pi^{\mathrm{A}} \cdot farms$	−0.76	1.69
$\pi^{\mathrm{A}} \cdot farms^2$	6.54[a]	1.53
P^{D}	1.47E–04	4.66E–04
$P^{\mathrm{D}} \cdot cpopd$	−1.03E–04	8.43E–05
$P^{\mathrm{D}} \cdot cpopd^2$	−6.46E–06[a]	2.80E–06
$P^{\mathrm{D}} \cdot vpopd$	1.26E–05[a]	1.45E–06
$P^{\mathrm{D}} \cdot vpopd^2$	−2.02E−11[a]	3.08E−12
$P^{\mathrm{D}} \cdot roads$	0.01[a]	3.00E–03
$P^{\mathrm{D}} \cdot roads^2$	−0.01[a]	4.74E–03
$P^{\mathrm{D}} \cdot farms$	−7.87E–03[a]	2.24E–03
$P^{\mathrm{D}} \cdot farms^2$	3.90E–03	2.29E–03
Connecticut	5304.86[a]	1905.22
Massachusetts	1975.95[a]	706.19
New Jersey	5406.47[a]	1336.2

$N = 2955$, $\overline{R}^2 = 0.67$. *cpopd* is the change in population density, *vpopd* is the variance of changes in population density, *roads* is highway density, *farms* is farmland density, π^{A} is the annual net return to agriculture, and P^{D} is the price of recently developed land.

[a] The estimate is significantly different from zero at the 5% level.

\$0.005. This result is unexpected, as Eq. (10) indicates that $\partial \overline{P}^{\mathrm{A}}/\partial P^{\mathrm{D}}$ should be positive. We can explain this finding by examining the estimates for individual counties. For counties near urban centers, the estimated values of $\partial \overline{P}^{\mathrm{A}}/\partial P^{\mathrm{D}}$ are almost always positive. Counties for which the estimates are positive (22% of all counties) have, on average, a population density almost three times that of counties with negative estimates. For the latter group of counties, the estimates of $\partial \overline{P}^{\mathrm{A}}/\partial P^{\mathrm{D}}$ are close to zero, reflecting the fact that land development is too

Table 2
The effects of the independent variables on the agricultural land value

Variable	Estimate	Standard error
π^{A}	5.00[a]	0.56
P^{D}	−0.005[a]	0.001
cpopd	65.14[a]	4.49
vpopd	0.45[a]	0.06
roads	1263.83[a]	101.56
farms	−390.77[a]	67.24

cpopd is the change in population density, *vpopd* is the variance of changes in population density, *roads* is highway density, *farms* is farmland density, π^{A} is the annual net return to agriculture, and $\overline{P}^{\mathrm{D}}$ is the price of recently developed land.

[a] The estimate is significantly different from zero at the 5% level.

far in the future to have much impact on agricultural land values. The measured effect for the average US county is correspondingly small (in absolute value) and negative. However, the average effect is positive when each estimate is weighted by the population density of the county.

A one unit increase in the rate of change in population density (*cpopd*) increases the average land value by \$65.14 per acre.[23] The variance of changes in population density (*vpopd*) is also found to have a positive effect on the current value of agricultural land. Consistent with the underlying theory, this finding suggests that option values associated with irreversible land development are capitalized into land prices. A one unit increase in highway density (*roads*) increases the average value of agricultural land by \$1264 per acre.[24] Higher highway density improves access to rural areas and should, therefore, increase the average value of agricultural land for development. Lastly, the share of the county land base in farmland has a negative effect on the average agricultural land value.[25] All else equal, more farmland dilutes the effect of higher future rents from development on the average value of agricultural land.

For comparison with the results discussed above, we also estimate a restricted version of the land price model in which the βs in (11) are constant across counties. In this case, (11) becomes $\overline{P}_i^{\mathrm{A}} = \beta_0 + \beta_1\pi_i^{\mathrm{A}} + \beta_2 P_i^{\mathrm{D}} + u_i$. Least squares results are $\hat{\beta}_0 = 642.67(27.33)$, $\hat{\beta}_1 = 6.51(0.17)$, and $\hat{\beta}_2 = 0.0064(0.0003)$,

[23] The average county in the continental US had a population in 1997 of approximately 80 thousand people and is roughly 600 thousand acres in size. Our results indicate that if the county's population were expected to increase by an additional 0.75% (600 people) per year in perpetuity, the average per-acre price of agricultural land would rise by \$65 today.

[24] In the average US county, this amounts to adding 600 miles of interstate highway or increasing the highway mileage by close to a factor of 10.

[25] In the average US county, a one percentage point increase in the farmland share reduces the average agricultural land value by \$3.91.

where standard errors are in parentheses, and the adjusted R^2 statistic equals 45%. The small coefficient on P_i^{D} indicates that, at the mean of the data, the development component contributes relatively little to the current price of agricultural land. This result is consistent with the estimated effects of P_i^{D} on land prices in the more general model. Note that the restricted model is nested in the more general model and is obtained through the following linear restrictions on the quadratic functions in (13): $c_{jk} = 0$ for $j = 0, 1, 2$ and $k = 1, \ldots, 8$. An F-test reveals that these restrictions are rejected at the 1% level,[26] indicating that the βs in (11) are not constant across counties and that the proxy variables in (13) are important determinants of land prices.

A primary goal is to compute an estimate of the agriculture and development components of the current value of agricultural land. These are given by respectively $\hat{\beta}_{1i}\pi_i^{\mathrm{A}}$ and $\hat{\beta}_{0i} + \hat{\beta}_{2i} P_i^{\mathrm{D}}$, where the hats indicate parameter estimates.[27] The results are summarized in Table 3 where we report the total current value of agricultural land for each state and the agriculture and development components of this value.[28] States are ranked according to the development component's share of the total current value. Northeastern states with large cities and little agricultural land are at the top of the list. For example, we estimate that in New Jersey approximately 80% of the value of agricultural land is attributable to future development rents. Some rapidly growing southeastern states (Florida, Tennessee, the Carolinas, Georgia) also show large values. California is relatively far down the list (number 30). Some counties in California have very high development shares, but most of the agricultural land is in the Central Valley region, relatively far from urban centers. Even so, the value of future land development capitalized into agricultural land values is \$5.8 billion in California, second only to Florida at \$8.7 billion. The value of future development on agricultural land is high in Illinois (\$1.8 billion), but this value is small compared with the total value of agricultural land in the state (\$57 billion), and Illinois is ranked near the bottom. For the contiguous US, we estimate the present value of future development on agricultural

[26] The test statistic is $(R_u^2 - R_r^2)(N-k)/(1-R_u^2)J$ where R_u^2 and R_r^2 are the R^2 statistics from the unrestricted and restricted least squares regressions, respectively, $J = 24$ is the number of restrictions, and $N - k = 2928$ is the degrees of freedom in the unrestricted regression. The test statistic equals 94.11, which exceeds the 1% critical $F_{J,N-k}$ value of 1.70.

[27] In computing these measures, one faces the somewhat arbitrary choice of including the intercept term in the agriculture component or in the development component. We include it in the agriculture component since to do otherwise would yield implausibly large development components for many rural and heavily agricultural counties. The intercept shifters for Connecticut, Massachusetts, and New Jersey are included with the development component because, as discussed above, it is likely that they capture unmodeled factors influencing future development rents.

[28] As discussed above, the sign of $\partial \overline{P}^{\mathrm{A}}/\partial P^{\mathrm{D}}$ is negative for some counties, which, in some cases, results in negative estimates of the development component. In computing the figures in Table 3, we set negative estimates of the development component equal to zero.

Table 3
The contribution of agricultural and future development rents to the 1997 value of US agricultural land, by state

State	Current value of agricultural land (million $)	Agriculture component (million $)	Development component (million $)	Development share of land value (percent)
NJ	5430	974	4457	0.82
CT	2126	414	1712	0.81
MA	2697	944	1753	0.65
FL	21,928	13,198	8730	0.40
NH	941	657	285	0.30
DE	1535	1072	463	0.30
MD	6798	4812	1986	0.29
SC	6871	5172	1699	0.25
PA	17,039	12,867	4172	0.24
NC	18,915	15,277	3637	0.19
TN	20,076	16,234	3842	0.19
RI	275	223	52	0.19
NY	9214	7561	1653	0.18
AL	12,530	10,376	2154	0.17
GA	15,987	13,349	2638	0.17
VA	15,606	13,062	2544	0.16
MI	16,433	13,792	2641	0.16
ME	1420	1201	219	0.15
VT	1914	1630	284	0.15
WV	3682	3188	494	0.13
AZ	8980	7848	1131	0.13
WI	18,561	16,306	2254	0.12
OH	28,791	25,601	3190	0.11
MS	10,645	9509	1136	0.11
OR	16,747	15,002	1745	0.10
LA	9454	8508	946	0.10
NV	1727	1566	162	0.09
UT	6887	6306	581	0.08
WA	18,189	16,676	1514	0.08
CA	72,570	66,767	5802	0.08
IN	31,225	28,810	2415	0.08
KY	19,311	17,982	1382	0.07
AR	16,616	15,570	1046	0.06
TX	77,373	72,758	4615	0.06
MO	30,837	29,159	1679	0.05
CO	19,849	18,884	965	0.05
ID	11,989	11,409	579	0.05
MN	30,285	29,141	1144	0.04
IL	57,031	55,219	1812	0.03
OK	20,250	19,728	522	0.03
NM	8473	8287	186	0.02
KS	26,655	26,185	471	0.02
MT	17,234	17,042	192	0.01
NE	29,599	29,305	295	0.01
IA	52,941	52,530	411	0.01
WY	7577	7528	50	0.01
SD	15,445	15,408	36	0.00
ND	15,801	15,801	0	0.00
US	863,352	780,785	81,699	0.09

land at $82 billion, which represents about 10% of the total value of agricultural land.

The results in Table 3 suggest that the influence of future land development on current land values depends jointly on the presence of urban areas and the current amount of agricultural land. This dependence is reinforced by examining the development component's share of the current land value for individual counties (Fig. 1). Future development rents are a relatively large component of agricultural land values along the west coast and in a large portion of the country east of the Mississippi River. The location of major urban centers (for example, Seattle, Denver, Minneapolis, the Boston–Washington corridor) are clearly seen. All of these counties are near or contain urban areas, have relatively little agricultural land, or both. In the Plains states from the Dakotas to Texas and in other heavily agricultural or rural states (for example, Iowa, Wyoming), future development rents contribute relatively little to average agricultural land values. In these cases, there is a large amount of agricultural land and little influence from urban areas.

4. Discussion and conclusions

We have conducted a national-level analysis of the determinants of agricultural land values to better understand how current land values are influenced by the potential for future land development. Our study makes two important contributions. First, we provide, to our knowledge, the first empirical evidence of the influence of option values on farmland values. In the theoretical model underlying our empirical analysis, option values arise from the stochasticity of future rents from land development and the irreversibility of land conversion. To capture the effects of uncertainty, we include a variable in the econometric model measuring the variance of annual changes in population density. The marginal effect of the population change variance on farmland values is positive and significantly different from zero, suggesting that option values associated with delaying irreversible land development are capitalized into the value of agricultural land. Option values have been shown to influence private land-use decisions (for example, Schatzki [28], Cho et al. [9]), but have not been considered in analyzes of farmland values.

A second contribution of this study is the decomposition of agricultural land values into components reflecting the discounted value of agricultural production and the discounted value of future land development. By identifying these price components, we can determine if landowners in a county face strong economic incentives to convert agricultural land. Previous studies have not yielded firm results on the magnitude of land development pressures due to their inability to separate the contributions of agricultural and development rent streams to the current price. Figure 1 reveals that future development rents are a substantial share of agricultural land values in areas surrounding urban centers. More generally,

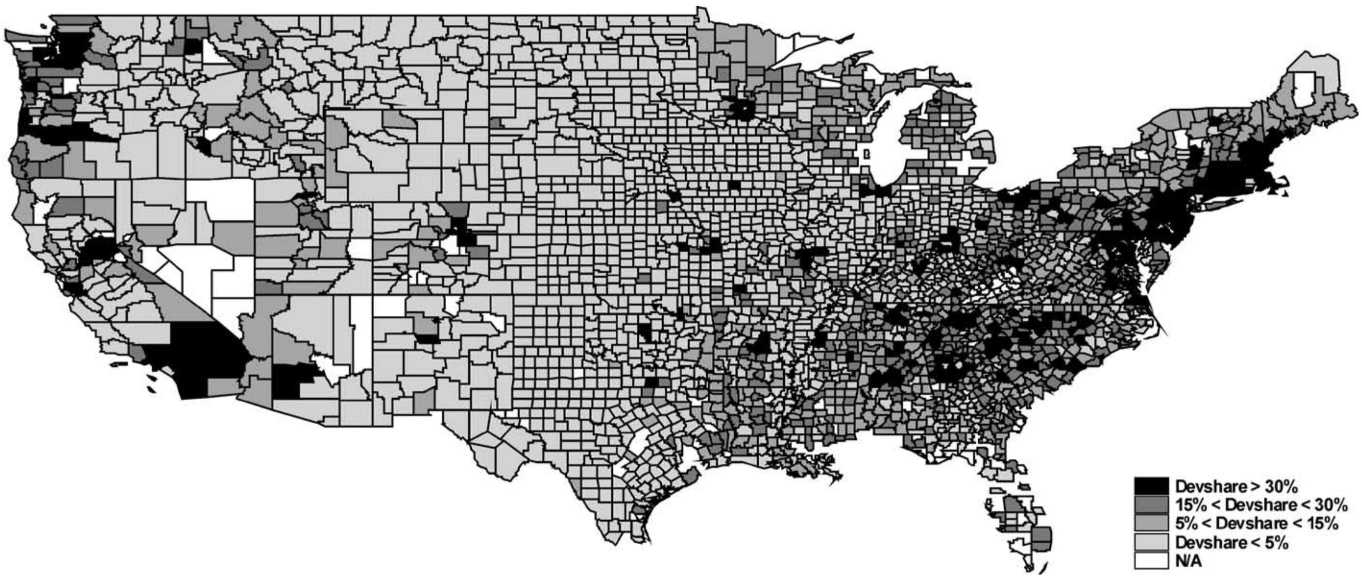

Fig. 1. The share of the 1997 value of agricultural land attributable to future development potential (devshare), by county.

relatively large development components are estimated for many counties east of the Mississippi River. Large development components can arise from strong pressures for land conversion, small amounts of agricultural land within the county, or some combination of both.

Our results on the contribution of future development rents to current agricultural land values yield a number of insights about policies to deter the conversion of agricultural land. As noted above, there has long been concern that the loss of productive agricultural land would substantially diminish the United States' capacity to produce food, with national as well as international consequences. Our results suggest that land development poses limited threats to food supply. We find that future rents from land development account for only about 10% of the current value of US agricultural land. Moreover, in most counties, including those in productive agricultural regions such as the midwestern US and the Central Valley of California, the development share of the current land value is typically below 5%. Thus, the evidence we obtain from decomposing agricultural land values does not suggest that large-scale development of the nation's productive agricultural lands is likely to happen soon. In part, this result reflects the relative abundance of land devoted to agricultural uses. For example, in many Iowa counties, over 90% of the land is in agriculture (statewide, the figure is 87%). In such cases, rents from future development, even if quite high, are effectively spread over many acres of land and, as a result, there is little effect on the average price of agricultural land.[29]

Even if loss of agricultural land is not a serious national security problem, it may have important consequences on a local level. Most states assess property taxes for agricultural land on the basis of value for agricultural production (Aiken [1]), but numerous studies have shown these programs to be ineffective at retaining agricultural land in rapidly developing areas (Malme [19]). Our results indicate that in counties near urban centers, future development rents often account for more than half of agricultural land values, suggesting that landowners would require substantial financial compensation to forego such development. Significant policies providing for the purchase of land or development rights will likely be required in these cases. By decomposing land values into agriculture and development components, we identify those counties where high land prices result from pressure for land development and, thus, where efforts might be directed to deter what are determined independently to be socially undesirable losses of agricultural land.

While our analysis yields a more complete description of the dynamic structure of agricultural land prices, it also raises issues that need to be addressed through further research. First, we provide evidence that farmland values are influenced by

[29] Fischel [13] observes that historical increases in urban land area are small relative to the total area of agricultural land, and reaches a similar conclusion regarding the threats posed by agricultural land development.

uncertainty over future development rents, but we do not know the magnitude of this effect. Thus, a topic for future research is the quantification of the option value's contribution to the current land price. Second, while we quantify the contribution of future development rents to the current land value, it is not entirely clear what this implies for the timing of land conversion. Use of the agricultural component of the land value to compute an implicit development time (from above, n) does not yield an estimate of the average conversion time for parcels within the county $(\bar{n})$. A topic left for future research is the recovery of the distribution of optimal development times for agricultural parcels within a county.

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Appendix A. Variable definitions and data sources

$\overline{P}_i^{\mathrm{A}}$ is the average price (dollars per acre) of agricultural land in county i in 1997. These data are reported in the Census of Agriculture and constructed as an average of owner-reported estimates of the current sales price of their farmland. The Census of Agriculture reports only the county average value. Data on individual owners are not disclosed for confidentiality reasons.

P_i^{D} is the average price (dollars per acre) of recently developed land in county i in 1997. P_i^{D} is estimated by backing out the average lot price from data on single-family home prices, which reflect both the value of structures and the land. Median prices for single family homes in 1980 and 1990 are taken from the decennial Census of Population and Housing Public Use Microdata Samples (PUMS 5% sample). This provides owner estimates of the market price of single-family homes at the level of county groups and subgroups. We consider only the value of single-family houses built within the five years preceding each census to ensure that the prices reflect the characteristics of the lots being developed and the houses being built in 1980 and 1990. Using 1980 and 1990 as base years, we extrapolate yearly data for each year between 1980 and 1997 using the Office of Federal Housing Enterprise Oversight (OFHEO) House Price Index. This index is based upon repeat home sales data and tracks quarterly changes in the price of a single-family home for each US state. While this data only provides the state average home price trend, we capture some of the county-level differences in annual home price changes by scaling the state trend up or down for each county to fit the change in home prices between 1980 and 1990 from the census. To back out the underlying land price for 1997, we multiply our annual estimate of the median single-family home price in each county by an estimate of the median share that the value of the lot represents in the total price of a single-family home. We compute this "lot share" from data in the annual Characteristics of New Housing Reports (C-25 series) from Census Bureau and the US Department of Housing and Urban Development. To obtain a per-acre measure of developed lot values, we divide the estimated median lot prices in each county by an estimate of lot sizes derived from the C-25 reports (making the assumption of constant returns to scale in land).

π_i^{A} is the average return (dollars per acre) to agriculture in county i in 1997. Using Census of Agriculture data, π_i^{A} is computed as $(TR_i - TC_i + GP_i)/A_i$ where TR_i is the value of all agricultural products sold, TC_i is total farm production expenses, GP_i are total government payments received by farmers, and A_i is total farmland area.

$cpopd_i$ is the average annual change in the total population of county i between 1990 and 1997, normalized on total county land area (in people per 1000 acres). Data are taken from the Census of Population.

$vpopd_i$ is the variance of annual changes in total county population over the period 1990 to 1997, normalized on total county land area (in people per 1000 acres).

$roads_i$ is the mileage of interstate and other principal arterial roads (for example, state highways) divided by total county land area (in highway miles per 1000 acres). Data were obtained from the US Department of Transportation.

$farms_i$ is measured as total farmland acres in 1997 divided by the county land area. Data are from the Census of Agriculture.

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