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Selected Paper

Location, Land Quality, and Rental Volatility

by

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Abstract: It appears to be widely believed that returns on low quality land are more variable than on high quality land. Using Ricardian rent as the measure of returns and sensitivity to output price as the measure of volatility, we investigate this null hypothesis for three different measures of quality. These are proximity to market, output productivity, and cost efficiency. In all cases, we identify precise conditions on the production technology such that rental volatility varies in a monotone manner with land quality. A method of econometric investigation of the relationship between rental volatility and land quality is developed and applied to Iowa cash rents data collected during 1994-2000. Our preliminary findings provide partial empirical support for the null hypothesis of an inverse relationship between quality and rental volatility with respect to commodity prices.

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Location, Land Quality and Rental Volatility

Introduction

As an input in production, land has value to the extent that it contributes to generating economic surplus. Thus, however measured, land value should bear an ordinal and monotone relationship with the rent it commands in the marketplace. But the relationship between land value and the variability of returns as product prices change is not immediately clear. The conventional approach to assessing land values, the Present-Value Constant Discount Rate (PV-CDR) model, does not appear to explain land prices very well. Falk (1991), Clark, Fulton and Scott (1993), Lence and Miller (1999), and others present evidence to this effect. But Lence and Miller cannot rule out that model when transaction costs are recognized and the planned holding time is short. Chavas and Thomas (1999) concur that transaction costs are important, and also find econometric evidence suggesting the importance of risk aversion as a determinant.

The focus of our attention is not, however, on the asset market but rather on the rental market where it is more likely that optimal pricing behavior occurs, where risk issues are of less consequence, and where transaction costs issues are not as severe. We seek to understand the relationship between land quality and the variability of rental returns as product prices change. The problem merits consideration because if a variant of the present value approach does underpin land pricing behavior then the relationship between land price volatility and land quality will work through the way in which land quality affects rental volatility. The problem is also important because commodity prices may remain depressed for sustained periods. And so the recipients of land rents most affected by depressed prices may face severe liquidity problems. Because there tends to be a strong positive spatial correlation between land quality attributes, the impact of depressed product prices may be systemic at the community and regional levels. This should be of concern to central bankers, especially in countries where asset and risk management markets are not well developed.

The effect of land quality on rent sensitivity to price fluctuations may also be important in environmental policy. The United States, among other countries, has periodically intervened in land rental markets in order to manage commodity supplies and to influence environmental conditions. The approach has been to offer long-term fixed rent contracts. The quality profile of land entering the program will depend upon the variability of rental returns on the land. One might expect that, due to risk aversion, if there is a strong positive correlation between rent variability and land quality then, other things fixed, the better quality land will be signed up. And conversely if the correlation is negative.

But little in the way of formal analysis has been conducted on the topic. The preponderance of opinion is that returns on lower quality land are more variable. Benirschka and Binkley (1994) identify anecdotal evidence to support this hypothesis. In their own empirical analysis they work with county-level land price data in Ohio, Indiana, Illinois, Missouri, and Iowa over the time intervals 1969-82 and 1982-87. Using loan rate data as a proxy for distance to market, and controlling for non-location determinants of quality, they find support for the hypothesis that land prices are more volatile in remote locations. But, while they suggest that the relation is an outcome of economic theory, they do not provide a detailed theoretical foundation for their findings.

Working at the rental level, the intent of this paper is to identify conditions on the production environment such that land rental value sensitivity to the output price environment is ordered in a monotone manner by land quality. We use three separate measures of quality; a geographical (i.e., basis) concept, a productivity index concept, and a cost penalty concept. In all cases we find that there is no theoretical reason why rental volatility should be inversely correlated with land quality. In the second part of the paper, we conduct an empirical investigation of the relationship between rental sensitivity to output price and land quality for 94 Iowa counties from 1994 to 2000. We conclude with some discussion of our estimation results.

Basis and Rent Volatility

A farm operator is assumed to cash rent cropland on an annual basis. The rent price is set in a competitive market at time 0, the beginning of the year, and harvest occurs at time 1, i.e., at year end. In seeking to establish the Ricardian rent, the operator establishes land profitability with reference to a locked in harvest-time futures price, F. The local basis at maturity is B and there is assumed to be no basis risk. Therefore, the operator can be assumed to be able to lock into a local price, through forward or futures contract positions, of $F + B$. Ignoring yield risk, the operator may ascertain that the maximized economic profit, gross of cash rent, is $R(F + B)$. In a competitive land market, this is the price that will pertain. From standard economic theory we know that this function is increasing and convex. We also assume that it is strictly positive. Our concern is with the volatility of this rental price. Denote the fractional sensitivity, or volatility, of the rent function by $V(F + B) = R'(F + B)/R(F + B) = d\ln[R(F + B)]/\mathrm{d}F$ where the number of primes indicates the order of differentiation.

Now, all other things fixed, let us move closer to the delivery point so that the basis increases toward zero. We seek to understand how $V(F + B)$ change with B. It is clear that the answer lies in the sign of

$$L(B) = \frac{dV(F+B)}{dB} = \frac{d^2 \ln[R(F+B)]}{dB^2} \qquad (1)$$

If $L(B) \leq 0$, then $R(F+B)$ is said to be log-concave over the domain. In this case the disadvantages in location exacerbate land rental volatility. If $L(B) \geq 0$ over the domain then $R(F+B)$ is said to be log-convex, and land rental volatility is ameliorated by remoteness from market.[1]

RESULT 1. *An increase in basis B reduces (increases) Ricardian rent volatility if the Ricardian rent function is log-concave (log-convex).*

Clearly, $L(B) \leq (\geq) 0$ according as

$$R(F+B)\,R''(F+B) \leq (\geq) [R'(F+B)]^2 \tag{2}$$

Note that concavity of $R(F+B)$ ensures log-concavity, but the converse is not true. We know that $R''(F+B) \geq 0$, so the sign of $L(B)$ cannot be immediately ascertained. Through a sequence of examples, we will show that it can plausibly be of either sign.

EXAMPLE 1: Suppose, as Benirschka and Binkley (1994) do in their graphical example, that the rental function is quadratic so that the supply function is linear. With $B = 0$, for convenience, write $R(F) = a_0 + a_1 F + a_2 F^2$, $a_0 < 0, a_1 \geq 0,\ a_2 > 0$. Then $L = [2a_0 a_2 - a_2^2 F^2 - (a_1 + a_2 F)^2] / R(F)^2 < 0$.

EXAMPLE 2: Suppose that the rental function is the power function, $R(F) = a_0 F^e$, $a_0 > 0, e > 1$. Then $L = -a_0^2\, e\, F^{2e-2} / [a_0^2 F^{2e}] < 0$.

[1] It is interesting to note that the log-concavity property has been widely studied in the statistics literature. See Shaked and Shanthikumar (1994) for statistical applications, while An (1998) and Chen (1997) provide economic applications.

EXAMPLE 3: Suppose that the rental function has expo-power function form, otherwise known as the Weibull functional form,[2] $R(F) = a_0 + a_1 \exp[a_2 F^a]$, $a_0 < 0$, $a_1 > 0$, $a_2 > 0$, $a \geq 1$. Here the coefficients have been restricted to ensure convexity and also that $R(0) < 0$. Then L has the sign of $a_0(a-1) + a_0\, a\, a_2 F^a + (a-1)a_1 \exp[a_2 F^a]$. It is immediate that there exists an interval $(a_0^*, 0]$ of strictly positive length such that $a_0 \in (a_0^*, 0]$ implies that $L > 0$.

This last example shows that theoretical support for the claim $L \leq 0$ must presuppose more structure on the profit function than just convexity.

Yield Productivity Index and Rent Volatility

Suppose that the quality of cash rented land could be captured by an exogenous parameter that multiplies a common production function. For a reference acre of unit quality, let input vector x permit production of up to $G(x)$. For an acre of quality $u > 0$, production is $u\,G(x)$. Suppressing factor prices, we can conclude that the Ricardian rent is $R(uF)$. For $u_2 > u_1$, an increase in land quality decreases (increases) rent volatility if $d \ln[R(u_2 F)] / dF \leq (\geq)$ $d \ln[R(u_1 F)] / dF$ on the relevant domain of F.

RESULT 2. *An increase in productivity index u reduces (increases) Ricardian rent volatility if the Ricardian rent function satisfies* $d^2 \ln[R(uF)] / dFdu \leq (\geq) 0$.

We have then that volatility increases for all $u > 0$ if

$$R(z)\, R'(z) + \{R(z)\, R''(z) - [R'(z)]^2\}\, z \geq 0\,,\ z = u\,F\ . \tag{3}$$

[2] See Saha (1993) for an application of this functional form as an utility function.

This is clearly a weaker condition than the log-convexity condition (2), and so example 3 is a case where an increase in quality increases rent volatility. Or for example 2, the power profit function, condition (3) is always satisfied with equality. And for example 1, condition (3) reduces to $a_0 a_1 + 4a_0 a_2 F + a_1 a_2 F^2 \geq 0$, a plausible requirement when F is sufficiently large.

Cost Efficiency Index and Rent Volatility

An alternative means of characterizing productivity is to identify the minimum cost of producing a given level of output. For baseline cost function $C(Q)$, consider a trajectory of augmented cost functions $C(Q) + l\, k(Q), l \geq 0, k(Q) > 0$. Ignoring basis, the rent function is $R(F, l) = \mathrm{Max}_Q\, F\, Q - C(Q) - l\, k(\mathrm{Q})$. Assuming that all cost functions are convex and that an interior optimum exists, denote the optimizing argument by Q^*. Our interest is in the sign of $d^2 \ln[R(F, ?)] / dFd\, ?$, i.e., of $\mathrm{T} = R(F, l)\, R_{F\,l}(F, l) - R_F(F, l)\, R_l(F, l)$ where subscripts are derivatives that have the obvious meanings. If the sign is positive (negative) then land that is more costly to work has a Ricardian rent function that is more (less) output price sensitive. Now from the envelop theorem we have $R_F(F, l) = Q^*$ and $R_l(F, l) = -k(Q^*)$. And a second derivative then gives $R_{F\,l}(F, l) = -k_Q(Q^*) / [C_{QQ}(Q^*) + l\, k_{QQ}(Q^*)]$. Now we evaluate all terms at $l = 0$ so that $F = C_Q(Q^*)$, and then

$$\mathrm{T} = -\frac{[Q^* C_Q(Q^*) - C(Q^*)]k_Q(Q^*)}{C_{QQ}(Q^*)} + Q^*\, k(Q^*).$$

Rearranging, and studying all values of Q, we see that the sign of T is that of

$$\frac{d\ln[QC_Q(Q) - C(Q)]}{dQ} - \frac{d\ln[k(Q)]}{dQ}.$$

In summary, we can conclude

RESULT 3. *An increase in cost index* l *increases* (*reduces*) *Ricardian rent volatility if* $d\ln k(Q)/dQ \leq (\geq) d\ln[QC_Q(Q) - C(Q)]/dQ$.

It is straightforward to show that for any twice continuously differentiable increasing and convex cost function there exists a $k(Q)$ such that the inequality holds in the forward direction and also a $k(Q)$ such that the inequality holds in the reverse direction. As to intuition for the result, note that an increase in l causes costs to increase because $k(Q) > 0$. Therefore $R(F)$ decreases. Now if $d\ln k(Q)/dQ$ is negative then an increase in l causes marginal costs to decrease because $k_Q(Q) \leq 0$. So the optimal level of output rises, i.e., $R_F(F)$ increases. Therefore the ratio $R_F(F)/R(F)$ increases.

Data

To provide some empirical evidence of how the sensitivity of rental rates to the output price relates to land quality, we studied the survey data *Cash Rental Rates for Iowa* collected during 1994-2000 (Edwards and Smith). These publications provide a reasonably accurate measure of typical cash rents for corn and soybeans on a county level. The cash rental rate in each county is typically calculated as an average of 12-18 responses of farmers, landowners, agricultural lenders, real estate brokers, and professional farm managers. There are at least five methods in setting cash rental rates: percent of land value, crop share equivalent, share of gross income, tenant's residual, and yield potential (Eggers, Mayer, and Thomas- EMT). In addition, cash rental arrangements can be classified as fixed or flexible depending on whether the amount of the rent is adjusted for actual yields and/or output prices (Langemeier). Also, payment provision may vary from one single annual payment to two installments (say a half is to be paid in March and the other half in fall). In the surveys done by Edward and Smith respondents supplied their information based on their best judgment assuming that half the rent is paid before planting and the remainder following harvest. Therefore, the recorded cash rental rates are likely to be relatively free from being biased due to the use of a particular method.

We seek to investigate how cash rents behave as a function of land quality characteristics. To do so, we use two readily available measures of land quality and location: (i) corn suitability rating (CSR) index measuring the quality of the land, and (ii) county loan rates used as a proxy for the basis. Basis variability at the county level, generally, reflects the distance to market (or a river to ship the harvest.) These measures of land quality and location are briefly described below.

CSRs provide a relative ranking of all soils based on their potential to be utilized for intensive row crop production (Iowa Agriculture, p.9). The CSR is an index that can be used to rate one soil's potential yield against another over a period of time. This index considers average weather conditions as well as frequency of use of the soil for row crop production. The most productive soils have an index of 100 and soils in Iowa usually range between 62 and 88. The construction of the CSR index is based on soil quality rather than actual yields. This makes the CSR index attractive to our study, as it does not embody any effects of the profit maximizing behavior and immediately provides us with a yield productivity index. Other measures of soil quality such as land capability classes may serve the same purpose and can be used to explain the land values. However, they appear extraneous in the study of the cropland cash rental rates in

Iowa as most of the farmland belongs to a single class. Further details about the CSR index, other soil characteristics, or land capability classes can be found in Iowa Soil Properties and Interpretations Database Manual (Iowa Agriculture).

The county loan rate is derived from the national loan rate and a twelve-month average posted county price (US Department of Agriculture). Production data and distance to the terminal markets are also taken into account in the setting of the loan rate. National average feed grain marketing assistance loan rates are set annually based on historical prices, the stocks-to-use ratio for corn, and a statutory maximum level of corn. For example, the corn loan rate is set at 85 percent of the simple average price received by producers during the marketing years for the immediately preceding 5 crops, excluding the highest and lowest prices. The posted county price is a price formulated to reflect actual market conditions in the county, and it is based on terminal market prices. Importantly, price adjustments are made by county and reflect market and transportation factors. Therefore, we obtain a good measure of the basis by using the county loan rates.

Econometrics

It seems reasonable to conjecture that the cropland rental rates are determined by expected price of output, productivity (yields), transportation costs (basis), and production costs. Hence, one should include among the explanatory variables the following: futures prices for corn and soybeans as those are the two main crops in Iowa, expected yields for corn and soybeans, basis for corn and soybeans, and production costs. However, accounting separately for corn and soybeans appears to lead to multi-collinearity among the independent variables. To avoid this problem, an index of expected net revenue (ENR) per acre received by farmers was constructed. The ENR index is calculated according to the following formula:

$$ENR_{i,t} = \frac{1}{2}\sum_{j}(Futures\ Price_{t,j} - Basis_{i,t,j} - Production\ Costs_{t,j}) \cdot Yield_{i,t,j}\ , \qquad (4)$$

where $Yield_{i,t,j}$ - expected yield in county i in year t for crop j,[3] and $j = Corn, Soybeans$.

Futures prices were calculated as a monthly average of year long corn and soybeans futures contracts traded on the Chicago Board of Trade with the expiration date in September. For example, "$\text{Futures Price}_{1995,\text{Corn}}$" is determined using the corn futures contracts maturing in September of 1996 as they were priced in August of 1995. The data on typical production costs in Iowa (state-average) was obtained from Duffy and Smith (2000). Note that the futures prices and the production costs are measured in dollars per bushel, and the yields are measured in bushels per acre.[4]

This is an open question whether averaging the net expected revenues per acre from corn and soybeans crops is an adequate weighting scheme. Comparison of the ratios of farmland sown to corn and soybeans over the 5 year period from 1995 to 2000 legitimizes the use of a simple average (Iowa Agricultural Statistics). Also, the observations for each county should bear different weights because the popularity of cash rental arrangements varies across the state. According to one survey, 79% of the farm leases in Northern Iowa are cash leases (EMT). Yet, only 36% of the farm leases in Southwest Iowa and 50% of the leases in South Central Iowa are cash leases. For example, EMT explain it by a greater variability in production in Southwest and

[3] The expected yields in (4) are calculated by averaging over corn and soybeans yields in the previous 10 years.
[4] The estimates of the CSR indexes and corn and soybeans yields in Iowa counties used in this study were taken from Edwards and Smith.

South Central Iowa. Unfortunately, such data (say the number of acres cash rented in each county) consistently collected during the period under study is not easily available. So every county's cash rental rate has equal weight in our model.

As pointed out by Lence and Miller, the cash rent time series (they find this for cash rents adjusted for property taxes and inflation) contains a unit root and therefore should be modeled in first-differences.[5] The time series of cash rental rates used in this study is relatively short (7 years), and therefore, any test for random walk type behavior is likely to be inconclusive. Furthermore, on economic grounds, it is hard to imagine that the cash rent wanders "too far" away from the base level, particularly, in the short-run.[6]

Note that during the period from 1994 to 2000 the change in both land quality measures, CSR and basis, occurs mainly across space (counties) while the change in the commodity price takes place across time. However, the change in the ENR index which measures the net expected revenues per acre of farm land varies along both dimensions due to different corn and soybean yields and basis in each location. Hence, to explain the variability in the cash rents one needs to use panel data if all of the available information is to be utilized. The constructed data set contains observations for 94 counties[7] during the time period from 1994 to 2000.

The estimation of the regression equation appears to be non-trivial when the observations are likely to be simultaneously time-wise and spatially correlated.[8] This is a somewhat novel feature of this analysis. There are several reasons to conjecture that both patterns of correlation

[5] Lence and Miller arrived at AR(1) model of the first-differenced real cash rents minus property taxes.
[6] The estimation using the first-differenced series of cash rental rates produces results similar to the ones found in this paper.
[7] Data on 94 out of 99 counties in Iowa were available for this study.
[8] For a recent review and discussion of the techniques used to model spatial autocorrelation in economics, please see Dubin (1998).

among the regression errors play an important role. Being a contractual arrangement, cash rental agreements are liable to exhibit lagged behavior over time. In addition, time and space correlation can be justified on the grounds of factors omitted from the regression (unobservable variables), for example, the cost efficiency index defined above.[9] Also, the data may exhibit spatial autocorrelation due to sharing of the common sources of information shared by the farmers in contiguous counties, using neighboring cash rental rates to negotiate your own, etc.

The choice of an estimation technique accounting for such an error structure is not clear-cut. Below we discuss some of the possible approaches. Also, there are at least two different ways of finding out whether there is a statistically significant relationship between the land quality attributes and the rental price sensitivity. To obtain a direct approximation to the second-order derivatives in question, the following regression equation is estimated:

$$\begin{aligned} \ln R_{i,t} &= a_0 + a_1 \ln R_{i,t-1} + a_2 ENR_{i,t} + a_3 CSR_i + a_4 Basis_{i,t} \\ &+ a_5 P_t \cdot CSR_i + a_6 P_t \cdot Basis_{i,t} + e_{i,t} \end{aligned} \tag{5}$$

where $R_{i,t}$ is the cash rent that was paid in county i during year t, $ENR_{i,t}$ is the expected net revenue index defined above, CSR_i is the corn suitability rating index, $Basis_{i,t}$ is replaced by the county loan rate in the actual regression, and P_t is the average of the corn and soybeans futures prices net of production costs per bushel. The CSR index and the loan rates have remained relatively constant (with a few minor exceptions) in Iowa counties over the considered time period, and therefore, for modeling purposes can be viewed as time-invariant. The lagged dependent variable is included to explicitly account for the infrequency of renegotiation of cash

[9] Alas, we have not been able to find data on a county level that is exogenous in the sense of not being a consequence of farmer's decisions, and allow us to construct an adequate approximation to the cost efficiency index.

rental rates arrangements. Note that given the panel nature of the data, we are interested in the signs of the regression coefficients on cross-product terms: $a_4 = \partial^2 \ln(R)/[\partial P\, \partial CSR]$, and $a_5 = \partial^2 \ln(R)/[\partial P\, \partial Basis]$.[10]

Another approach is to create dummy variables corresponding to the quality classes as measured by the CSR index and the distance from the terminal market, then test for their joint and/or individual significance. Then the regression equation becomes:

$$\begin{aligned} \ln R_{i,t} &= b_0 + b_1 \ln R_{i,t-1} + b_2 DC_i + b_3 DB_i + b_4 DC_i \cdot DB_i + b_5 ENR_{i,t} + b_6 DC_i \cdot P_t \\ &+ b_7 DB_i \cdot P_t + b_8 DC_i \cdot DB_i \cdot P_t + e_{i,t} \end{aligned}, \qquad (6)$$

where $DC_i = \begin{cases} 1, if & CSR_i \text{ is high} \\ 0, if & CSR_i \text{ is low} \end{cases}$, $DB_i = \begin{cases} 1, if & \text{Basis}_i \text{ is high} \\ 0, if & \text{Basis}_i \text{ is low} \end{cases}$.

The formulas for the specified relationships between the cash rents and the output price depending on the class of the farm land are succinctly summarized in Table 1.

We now turn to the next issue of choosing an econometric technique to estimate these regression equations. As noted above, it is very likely that the errors exhibit both patterns of autocorrelation: across time and across space. Therefore, the following error structure is proposed:

$$e_{i,t} = r e_{i,t-1} + u_{i,t}, \text{ where } u_{i,t} = l \sum_{j=1}^{N} w_{ij} u_{j,t} + v_{i,t},\ v_{i,t} \sim i.i.d.\, N(0, s^2). \qquad (7)$$

Here r is interpreted as a time autocorrelation coefficient, l is a spacial autocorrelation coefficient, and w_{ij} is an element of a weighting contiguity matrix W, which is positive if

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[10] These are the "short-run" responses. The "long-run" measures of the sensitivity are given by $[1/(1-a_1)]\partial^2 \ln(R)/[\partial P\, \partial CSR]$ and $[1/(1-a_1)]\partial^2 \ln(R)/[\partial P\, \partial Basis]$.

counties *i* and *j* are contiguous and zero otherwise (for example, see Benirschka and Binkley (1994)). The rows of the weighting matrix are normalized so that the sum of the elements of each row is one. Rewriting (7) in a matrix form, we have:

$$e_t = re_{t-1} + u_t \text{, where } u_t = lWu_t + v_t,\ v_t \sim N(0, s^2 I). \tag{8}$$

Here e_t, u_t, v_t are Nx1 vectors, W is an NxN matrix. Hence, we can write $u_t = (I - lW)^{-1} v_t$.

An OLS-type procedure similar to Cochrane-Orcutt estimation procedure for AR(1) errors can be employed to obtain the estimates of both autocorrelation parameters. The algorithm may consist of the following steps: 1) use OLS to obtain estimated residuals; 2) use OLS residuals to get an estimate of the time autocorrelation parameter; 3) use it to transform the variables (partial differencing); 4) estimate the transformed model using OLS and obtain the residuals; 5) using OLS on these residuals obtain the estimate of spatial autocorrelation; 6) transform the variables and iterate steps 1-6 until convergence. However, the small sample properties of such estimators are not fully known. In addition, the usual caveats of using OLS in the presence of the lagged dependent variable and autocorrelated disturbances apply (Green, pp.586-590). In the absence of the lagged dependent variable, as in the case of the usual time autocorrelation, OLS estimates remain unbiased and consistent (Cliff and Ord). However, it is no longer efficient and OLS standard errors are likely to be estimated incorrectly (too small if the correlation coefficient is positive, and too big otherwise). The parameter estimates obtained using the algorithm presented above differ from the estimates produced by the maximum likelihood approach.

The parameters of the regression equation along with the correlation coefficients can be estimated using the maximum likelihood estimation (for example, econometric software GAUSS provides a convenient routine for maximizing a likelihood function.) For simplicity, we rewrite the regression equation in a general form $y_{i,t} = x'_{i,t} b + e_{i,t}$, or in vector form:

$$y_t = x_t' b + e_t, \tag{9}$$

where y_t, x_t, e_t are Nx1 column vectors containing observations for N counties in year t.

Transforming Eqn (9) through partial differencing, we obtain:

$$y_t^* = x_t^{*\prime} b + u_t, \tag{10}$$

where $y_1^* = \sqrt{1-r^2}\, y_1, x_1^* = \sqrt{1-r^2}\, x_1$ and $y_t^* = y_t - r y_{t-1},\ x_t^* = x_t - r x_{t-1},\ \forall t > 1$.

Substituting for u_t from (8), we have:

$$y_t^* = x_t^{*\prime} b + (I - lW)^{-1} v_t, \tag{11}$$

and transforming Eqn (11) by multiplying it through by $(I - lW)$:

$$y_t^{**} = x_t^{**\prime} b + v_t,$$

where $y_t^{**} = (I - lW) y_t^*,\ x_t^{**} = (I - lW) x_t^*$.

This equation can be directly estimated by maximizing the likelihood function obtained by applying the change of variable formula where necessary (note that only variable $y_{i,t}$ is directly observable):

$$L(Y) = (2p)^{-\frac{TN}{2}} \left(s^2\right)^{\frac{-TN}{2}} \left(\sqrt{1-r^2}\right)^{\frac{N}{2}} \left|I - lW\right|^T e^{-\frac{1}{2s^2}\sum_{t=1}^{T}(y_t^{**} - x_t^{**\prime} b)^2},$$

where $|I - lW|$ is the absolute value of the determinant of the transformation matrix.

Estimation Results

Taking up the first approach of directly estimating the effect of land quality and distance to market, we estimate Eqn (5) using OLS (taking no account of time and space autocorrelation among the residuals) and MLE procedures. The estimation results are presented in Table 2.

As mentioned above, the use of OLS is problematic when the lagged values of the dependent variable are included among the explanatory variables and errors are autocorrelated. The comparison of the OLS and ML estimated coefficients and their variances corroborates to the well-known property that the OLS estimators tend to be biased and inconsistent as well as inefficiently estimated in such cases. Therefore, we focus on the ML estimation results. All coefficients have the intuitively expected signs. The effect of the previous year cash rental rate is large and very significant. This agrees with the sluggish adjustment of farmland rental contracts. The expected revenue generated from an acre of cropland also seems to be an important determinant of the cash rental rate. Both spatial and time autocorrelation among the regression errors appear to be present and significant.

However, the effects of neither yield characteristics nor transportation costs are statistically significant under the imposed error correlation structure. As expected, an acre with a higher CSR fetches a higher rent. Whereas, the cash rent paid for a similar acre in a county that is more distant from the point of delivery tends to be lower. Also, the relationship between the sensitivity of the rental rate to the output price and the land quality seem to conform to the proposition suggested by Benirschka and Binkley. Namely, both a higher CSR and lower transportation costs (lower basis) imply a decline in the sensitivity of the rental rates to output price. Such an inverse relationship is somewhat more pronounced when the CSR index increases rather than when the basis falls.

A variety of hypotheses about the presence of any relationship between the rental rate sensitivity and land quality characteristics have been tested using likelihood ratio tests. The specified hypotheses include the following: $H_0 : a_5 = a_6 = 0$, $H_0 : a_5 = 0$, $H_0 : a_6 = 0$, $H_0 : a_5 = 0$ given $a_6 = 0$, and $H_0 : a_6 = 0$ given $a_5 = 0$. In words, we test for the presence of the CSR index and/or basis effect, the CSR (basis) effect while accounting for the basis (CSR), and the CSR (basis) effect given no basis (CSR) effect, respectively. Both individual and joint quality effects on rental rate sensitivity are rejected at the conventional levels of significance. In addition, we fail to reject the null hypothesis of no CSR (basis) effect either in the presence or lack of the basis (CSR) effect.

The results of estimating Eqn (6) using OLS and MLE are presented in Table 3. Again F-tests for joint significance and likelihood ratio tests testify to the lack of any impact of quality measures on rental price sensitivity. However, the adopted error structure is strongly supported by the data. Economically meaningful tests for the presence of the quality effects when dummy variables are included in the regression are essentially the same as before. To encompass all the possible combinations of interaction between the CSR and basis effects on the volatility, two more regressions are estimated: one that isolates the CSR effect by setting $b_2 = b_3 = b_7 = b_8 = 0$, and another one that isolates the basis effect by setting $b_1 = b_3 = b_6 = b_8 = 0$. The following set of hypotheses has been tested: $H_0 : b_1 = b_2 = b_3 = b_6 = b_7 = b_8 = 0$, $H_0 : b_1 = b_3 = b_6 = b_8 = 0$, $H_0 : b_2 = b_3 = b_7 = b_8 = 0$ $H_0 : b_1 = b_6 = 0$ given $b_2 = b_3 = b_7 = b_8 = 0$, and $H_0 : b_2 = b_7 = 0$ given $b_1 = b_3 = b_6 = b_8 = 0$.

In words, we test for the presence of either quality characteristic effect, then we restrict our attention to each effect individually, and finally, we investigate the effects in isolation. One might speculate testing for joint CSR and basis effect for different quality classes (high/low CSR, high/low Basis.) Such tests would be difficult to interpret and may be judged as “mining” the data. The likelihood ratio test results of the estimation of the individual and/or joint quality effects reveal that none of the hypotheses can be rejected at the conventional levels of significance. The conclusion drawn by both approaches is the apparent lack of any statistically significant relationship between the land quality and rental price volatility.

A somewhat interesting issue is the sign of the time autocorrelation coefficient. The value of r is fairly robust to the regression specification (which land quality attributes are included among explanatory variables) and error structure (whether the spatial autocorrelation parameter is restricted to zero.) As pointed out before, the presence of time autocorrelation can be interpreted either as a consequence of omitting factors pertinent to the time-series model explaining the cash rents, or a lag occurring in the landlord-tenant negotiation behavior. Observe that $Cov(y_{i,t-1}, e_{i,t})$ inherits the sign of time autocorrelation parameter r.[11] This implies a negative correlation between the previous year cash rental rate and the likely deviation from the stable relationship based on the relevant economic factors in the subsequent year.

Discussion

These empirical results provide limited support for the view prevailing in the literature that posits a negative relationship between the quality of land and the land price sensitivity to the price of output. As discussed above, there is no theoretical reason to presume a certain nature of the relationship unless one is willing to impose some structure on the underlying production

[11] For details, please see Green (1997, p. 587)

technology. Data for Iowa counties gathered during the period from 1994 to 2000 testifies to the lack of any statistically significant correlation between cash rent sensitivity to output price and the land quality.

There are several reasons to cast some doubt on the validity of these empirical findings. First of all, it is possible that the county level data does not provide sufficient variability in the explanatory variables. Typically, farmland in Iowa is of superior quality that is more or less stable over the state and the geographical locations of the counties are relatively immaterial with respect to the shipping expenses. Note that the variability in the loan rates across counties in Iowa, measured as a percentage of the output price, can be a small number (say, for the largest difference in the loan rates (basis) between the counties this number is 0.19/2.5 x 100% =7.6%). A partial remedy for that problem would be to analyze the data on the county level but for several corn and soybean producing states. Note that eliminating the time dimension from the analysis does not lead to a more statistically significant relationship. Studying the change in the cash rents only in the space dimension yields mixed results.

Another possible pitfall is the accepted error correlation structure and/or estimation technique. If the analysis is limited to only accounting for autocorrelation in one dimension (either time or spatial) then the statistical significance of the estimated parameters will be affected.[12] As it is well known in the literature, not allowing for a positive autocorrelation leads to underestimation of coefficient variances, while the opposite holds for negative autocorrelation. Hence, it is important that the imposed error structure does not bias the tests for statistical significance. In our case, the likelihood ratio tests for individual and joint significance of r and

[12] Also, see Benirschka and Binkley (1994) for a discussion of the impact of the arbitrariness of the county

l confirm the pertinence of the suggested pattern of time and space autocorrelation among the errors.

Concluding Remarks

The main point of this paper is the observation that, at least theoretically and in a world of optimizing agents, rental volatilities bear an indeterminate relationship with land quality. And so evidence on that relationship reveals something about the nature of the underlying technology. To be more specific on this point, observe all three of our results that concern the sign of $d^2 \ln[R(F,t)]/[\mathrm{d}F\,\mathrm{d}t]$ where t is one of $\{B, u, l\}$ and where additional structure might have been imposed on the rent function. When $d^2(\ln[R(F,t)]/dFdt \geq 0$ on the relevant domain, then the rent function is said to be log-supermodular. If the sign is negative on the relevant domain then the rent function is log-submodular. The log-supermodularity property has arisen in economic analyses of uncertainty by Athey (1996) and others.

Consider the implications of finding no significant quality effects for the functional form of Ricardian rent $R(u(F+B))$. The empirical analysis asserts that $L(x) = dV(x)/dx = 0$, where x represents either the futures price, or the basis, or the sum of the two (see Eqn (1)). Solving this differential equation we have $R(x) = c_2 e^{c_1 x}$, where c_1 and c_2 are arbitrary constants. This is the functional form of the relationship between the rental rate and the farm gate output price implied by the lack of the basis effect on rental volatility.

On the other hand, substituting $L(u(F+B)) = 0$ in $d^2 \ln[R(u(F+B))]/[d(F+B)du] = 0$, we obtain $R'(F+B) = 0$, which can be tested empirically. This hypothesis is immediately

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definition on the estimated variances.

rejected by the data. There is a very strong (and intuitively obvious) positive relationship between the expected output prices and the cash rent found in the data. Hence, we conclude that describing the productivity index as an exogenous parameter augmenting the output price by a multiplicative factor may not be appropriate since it has little empirical support. Of course, other ways of representing the yield productivity index are possible. Or, alternatively, the finding of no relationship between land quality and the rental rate sensitivity may be erroneous.

If the intent of an economic inquiry is to ascertain the role of asset quality on rent or asset price, then it would seem wise to pose a maintained production technology that is flexible in the sense that it admits log-supermodularity and log-submodularity with respect to the measure of quality. Otherwise the object of inquiry may be imposed by the technology.

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Table 1. Dummy variables Approach

		CSR	
		Low	High
Basis	Low	$b_0 + b_5 ENR_{i,t}$	$b_0 + b_2 + b_5 ENR_{i,t} + b_6 P_t$
	High	$b_0 + b_3 + b_5 ENR_{i,t} + b_7 P_t$	$b_0 + b_2 + b_3 + b_5 ENR_{i,t} + (b_6 + b_7 + b_8) P_t$

Table 2. Direct Estimation of Quality Effects on Rental Volatility

Parameter	OLS	MLE	Variable Description
a_0	1.1817*	0.9322*	Intercept
	(14.6602)	(11.080)	
a_1	0.6889*	0.7824*	Lagged log of Cash Rent ($/acre)
	(30.7300)	(33.109)	
a_2	0.0027*	0.0012*	ENR (Expected Net Revenue) ($/acre)
	(7.9378)	(3.181)	
a_3	0.0029*	0.0006	The CSR index (%)
	(3.9651)	(0.772)	
a_4	0.0727	-0.1328	Basis (county loan rate) ($)
	(0.4563)	(-0.740)	
a_5	-0.0015*	-0.0004	The CSR index times Net Price ($/bushel)
	(-4.7079)	(-1.019)	
a_6	-0.0058	0.0571	Basis times Net Price ($/bushel)
	(-0.0694)	(0.607)	
l	-	0.3998*	Spatial autocorrelation
		(7.903)	
r	-	-0.3348*	Time autocorrelation
		(-6.383)	

OLS: $R^2 = 0.88$, $\overline{R}^2 = 0.878$.

Note: * denotes significantly different from zero at the 5% (1%) level of statistical significance, numbers in parenthesis are t-statistics, 564 observations.

Table 3. Dummy Variables Approach Estimation Results

Parameter	OLS	MLE	Variable Description
b_0	1.1065*	0.9071*	Intercept
	(13.3838)	(10.116)	
b_1	0.7343*	0.7908*	Lagged log (Cash Rent) ($/acre)
	(36.0288)	(36.840)	
b_2	0.0843*	0.0046	Dummy for the CSR index
	(2.4996)	(0.146)	
b_3	0.0901*	0.0153	Dummy for Basis (county loan rate)
	(3.1389)	(0.518)	
b_4	-0.091609*	-0.0092	Dummy for CSR times Dummy for Basis
	(-2.228659)	(-0.247)	
b_5	0.0017*	0.0011*	ENR (Expected Net Revenue) ($/acre)
	(7.3199)	(4.668)	
b_6	-0.0508*	-0.0059	Net Price ($/bushel) times Dummy for CSR
	(-2.7043)	(-0.341)	
b_7	-0.0547*	-0.0138	Net Price ($/bushel) times Dummy for Basis
	(-3.3706)	(-0.823)	
b_8	0.0564*	0.0105	Net Price ($/bushel) times Dummy for CSR times Dummy for Basis
	(2.4827)	(0.509)	
l	-	0.4051*	Spatial autocorrelation
		(7.597)	
r	-	-0.3417*	Time autocorrelation
		(-7.046)	

OLS: $R^2 = 0.876$, $\overline{R}^2 = 0.875$.

Note: * denotes significantly different from zero at the 5% (1%) level of statistical significance, numbers in parenthesis are t-statistics, 564 observations.