Modeling Nonlinearities in Farmland Values: A Dynamic Panel Threshold Error-Correction Model

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Abstract

Earlier studies usually indicate that farmland prices and cash rents are not cointegrated, a finding that seems at odds with the implications of the present value model. The main objective of this study is to explore whether this absence of empirical support for the present value model can be attributed to the restrictiveness of conventional time series methods. I suggest a panel unit root model with two regimes in which the adjustment process may be characterized by the presence of thresholds and discontinuities reflecting the presence of transactions costs and other barriers to adjustment. Using farmland value and cash rents data for 10 agricultural states of the U.S. between 1960 and 2008, empirical findings give modest improvement over the linear unit root process. It is suggested that there might be a bias caused by cross sectional dependence and an inadequate time span of the data.

Keywords: Present Value Model; Transactions costs; Thresholds; Panel unit root

1. Introduction

Farm real estate is the major asset on the farm sector balance sheet, accounting for nearly 79 percent of total U.S. farm assets in 2000.[1] Boom-boost cycles in farmland prices trigger noticeable wealth changes in the farm sector as farmland is the most important asset in the sector. There has been a rapid increase in farmland values during the 1970s and early 1980s which was followed by a sharp decline during 1982-87. The slow upward trend in land prices beginning in 1987 began to accelerate in 1994. The value of U.S. farm real estate, including all agricultural land and buildings, averaged $2,350 per acre on January 1, 2008, 8.8 percent higher compared to 2007 and 290 percent higher compared to 1987 (USDA- ERS, Land Use, Value and Management Briefing Room) .

A common approach to valuing farmland is based on a net-present value model (PVM) in which the current value of an acre of land is modeled as the sum of expected future cash flows, discounted according to the risk of individual sources of these cash flows. The empirical model is based on the following assumption (Campbell and Shiller, 1988; Falk, 1991): The present value model of real land prices requires that, if the real rent possesses a unit root, then so must the price of land itself. Furthermore, assuming a fixed, constant discount rate, land prices and rents must be cointegrated with a unit coefficient on rents. However, there has been a divergence between farmland values and returns to land from agriculture. Market prices of land have increased significantly over the past decade, while there has been little increase in the cash flow generated from farmland. Falk (1991) analyzed Iowa farmland prices and ended up rejecting the PVM. Similar results were reported for Illinois by Clark et al. (1993). Tegene and Kuchler

[1] Other 21% consists of livestock and poultry, machinery and motor vehicles crops and financial assets.

(1993) and Engsted (1998) examined three US regions—the Lake States, the Corn Belt and the Northern Plains—and found no evidence supporting the present value model. The divergence between the present value of future cash flows and the market price of farmland suggests that other factors beyond returns from agriculture may play a role in determining land values.

One possible explanation for the lack of consensus about farmland pricing and the explanatory power of the PVM might be the presence of market frictions. Market frictions, including transactions costs, may drive a wedge between the price at which outsiders wish to buy land and that at which farmers wish to sell it. The market price can be anywhere within this wedge, and can easily deviate from its frictionless present value. One can interpret this wedge as a band of inaction inside which farmers neither buy nor sell land, even in the face of changing expected returns. Such a band would be centered on the price that would prevail in the absence of transaction costs, and its width would be determined by the size of these costs. Transaction costs and the large capital investments necessary to participate in the agricultural land market may cause nonlinearities in the adjustment of values and rents towards long run equilibrium. In particular, rental agreements in agricultural land markets may be relatively fixed in the short run and thus significant transactions costs may be associated with the renegotiation of rental contracts. Although the costs associated with trading many financial assets are small, costs involved in transferring ownership of farmland typically exceed 7-8% of the purchase price (De Fontnouvelle and Lence, 2002).

Just and Miranowski (1993) was the first study to incorporate explicitly the potentially large transaction costs involved in the transfer of ownership of farmland. They developed a structural model of farmland prices that explicitly accounted for a large number of relevant issues, such as the "multidimensional effects of inflation associated with capital erosion, savings-return erosion, and real debt reduction as well as the effect of changes in the opportunity cost of capital" (p. 168). However, they did not test specifically whether transaction costs are the reason behind the PVM failure. Chavas and Thomas (1999) introduced a dynamic theoretical model of land prices, allowing for nonadditive dynamic preferences and risk aversion, as well as transaction costs. Chavas and Thomas (1999) perform Generalized Method of Moments analysis of U.S. land values for the period 1950-96. Their findings indicate that transaction costs have significant effects on land prices.

Lence (2001) argues that both studies mentioned above fail to recognize the non-stationary nature of farmland prices. Lence and Miller (1999) analyzed farmland prices in the presence of proportional transaction costs while restricting their attention to the widely used constant-discount rate version of present value model (CDR-PVM). They reformulated the CDR-PVM accounting for the frictions created by transactions costs. Their empirical analysis is based on an autoregressive model of the stochastic discounted excess return. Using Iowa data for the 1900-1994 period, they find mixed evidence regarding the CDR-PVM in the presence of transaction costs. CDR-PVM was consistent with typical transaction costs assuming a one-period holding horizon, but not when an infinite-holding horizon was hypothesized. Defontnouvelle and Lence (2002) use kernel regressions to test the theroretical model of Lence and Miller (1999). They also expand Lence and Miller's(1999) data set to include 20 major agricultural states between 1921-1990, as well as two different national series. They confirm that the behavior of land values and rents is consistent with the CDR-PVM in the presence of typical transaction costs. Finally, Lence

(2003) tests the model in Lence and Miller (1999) using a threshold autoregressive model (TAR) rather than a standard one. Using Iowa farmland data over 1900-1994 period, Lence (2003) argues that TAR gives a better representation of farmland-pricing behavior. He finds that land price behavior is consistent with the necessary conditions for market equilibrium under rational expectations and the typical transactions costs observed in land markets.

Another explanation for the absence of empirical support for the present value model might be that standard tests may not be powerful enough to detect long run equilibrium when applied to single, short time series. A promising approach would be to combine the sample information from the time series dimension with that from the cross-section. Panel data methods are expected to be more accurate than conventional methods based on single time series. As to my knowledge, the only study using a panel data time series framework to test the Present Value Model of U.S. farmland is by Guierrez et. al. (2007). They argue that the failure to find cointegration between farmland prices and cash rents may be due to the low power of their tests and to the presence of structural change representing a shifting risk premium on farmland investments. They use panel unit root and cointegration that allow for breaks in the cointegration relationship. Empirical results, based on a panel covering 31 U.S. states between 1960 and 2000, suggest that the present value model of farmland prices cannot be rejected once accounted for the structural change in early 1980s.

The main objective of this study is to explore whether this absence of empirical support for the present value model can be attributed to the restrictiveness of conventional time series methods. Specifically, I consider a panel unit root model with two regimes in which the adjustment process may be characterized by the presence of thresholds and discontinuities reflecting the presence of transactions costs and other barriers to adjustment. In this case, the present value model may not imply stationarity in the usual sense, and the appropriate empirical test is therefore not a conventional unit root test, but rather a test for unit roots with thresholds or other discontinuities in adjustment. A threshold autoregressive panel unit root approach is proposed by extending the technique introduced by Caner and Hansen (2001) to a panel data context using Fisher-type tests as proposed by Maddala and Wu (1999) and Choi (2001).

The remainder of this paper is organized as follows. Model is presented in Section 2. Section 3 describes the methodology. In Section 4, these methods are applied to a panel of 10 agricultural states of the U.S., for which time series data on farmland prices and cash rents are available for 1960–2008. The results show that, although there is an improvement over the conventional unit root procedures, the present value model is still rejected after accounting for nonlinearities. Section 5 gives a discussion on results and a direction for further research.

2. Present Value Model

The analysis in this paper is based on Present Value Model assuming time-varying expected returns. It is more complicated compared to the case of a constant discount rate because the relation between prices and returns becomes non-linear. Campbell and Shiller (1988) propose a log-linear approximation of the present value framework which enables to investigate stock prices behavior under time-varying discount rates. The Campbell and Shiller (1988) version of the PVM relates, P_{it}, the real price per acre of farmland in state $i=1, \dots, N$ at period $t = 1, \dots,$

T, to C_{it+1}, the real rent per acre of farmland paid at beginning of time *t+1* for the land held from the beginning of time t to the beginning of time *t+1*. In this notation, the log of the gross real rate of return on an acre of land in state i from period t to *t+1*, R_{it+1} may be defined as:

$$\log(R_{i,t+1}) = \log(P_{i,t+1} + C_{i,t+1}) - \log(P_{i,t})$$

or equivalently;

$$r_{i,t+1} = \log(1 + \exp(s_{i,t+1})) + p_{i,t+1} - p_{i,t} \tag{1}$$

where the lowercase letters are the logs of the corresponding uppercase variables and $s_{i,t+1} = c_{i,t+1} - p_{i,t+1}$ is the log of the ratio of rents to prices, usually referred to as the 'spread' in the financial literature. The objective here is to write log returns as a linear function of log prices and log rents, which is complicated by the first term on the right-hand side of equation (1). Campbell and Shiller (1988) shows that, this problem can be solved by linearizing $\log(1 + \exp(s_{i,t+1}))$ around the time series mean of $s_{i,t+1}$, s_i. Equation (1) can be written approximately as:

$$r_{i,t+1} \approx k_i + s_{i,t} - \rho_i s_{i,t+1} + \Delta c_{i,t+1} \tag{2}$$

where $k_i = -\log(\rho_i) - (1 - \rho_i)\log(1/\rho_i - 1)$ and $\rho_i = 1/(1 + \exp(s_i))$ are parameters of the linearization.

Equation (2) is basically an ordinary difference equation, which can be solved forward indefinitely to obtain the following approximation for s_{it} :

$$s_{i,t} \approx -k_i / (1 - \rho_i) - \sum_{m=0}^{\infty} \rho_i^m (\Delta c_{i,t+1+m} - r_{i,t+1+m}) + \lim_{m \to \infty} \rho_i^m s_{i,t+m} \tag{3}$$

Taking expectations conditional on all the information available at time t and imposing the transversality condition (i.e., expected value of the last term on the right-hand side of equation (3) is zero) the above equation can be rewritten as

$$s_{i,t} \approx -k_i / (1 - \rho_i) + E_t \left[\sum_{m=0}^{\infty} \rho_i^m (-\Delta c_{i,t+1+m} + r_{i,t+1+m}) \right] \tag{4}$$

This equation expresses the current value of s_{it} in terms of the present discounted value of expected future values of Δc_{it+1+m} and r_{it+1}. Given that changes in the log cash rents and the log discount rate follow a stationary process, the log land price and the log cash rents are cointegrated with the cointegrating vector [1, -1] and the log rent-to-value ratio (spread) is a stationary process.

To understand this cointegration relation, one may intuitively think that if current land values are high in relation to current cash rents (i.e. investors are willing to pay more or the land is overpriced), cash rents are expected to grow. If agents are fully rational under the PVM, prices and cash rents cannot drift arbitrarily and persistently far apart and the ratio will show a reverting behavior towards an attractor.

A popular approach in testing the PVM is to test the cointegration between log prices and log cash rents series. However, cointegrating and error correction models are complicated when considering threshold behavior *and* panel data together. Therefore, this paper relies on univariate panel unit root procedures rather than multivariate cointegration procedures since the theoretical model at hand allows doing so (Equation (4) then implies that the stationarity of the spread can be viewed as a necessary condition for the validity of the present value model). In other words, the analysis is based on testing the stationarity of log rent-to-value ratio in a panel data framework. The empirical investigation of the log rent-to-value model, first, does not involve the estimation of an unknown cointegrating parameter and, second, measurement problems associated with deflating nominal stock prices and dividends by some price index do not occur. With the exception of highly persistent expected returns and small samples, cointegration tests on the log dividend–price ratio tend to reject the null hypothesis of no cointegration more frequently than cointegration tests in levels of price and dividend series (Bohl and Siklos, 2004). Working with the logs of rent and values instead of working with the levels is also consistent with relaxing the assumption of constant expected returns in favor of time-varying expected returns.

3. Methodology

Analysis has two stages. First, the unit root process in the logarithmic rent-to-value ratio is tested for each individual state and for the entire panel assuming a linear adjustment process. At the second stage, a threshold is assumed to be in effect in each of the farmland markets, where the transaction costs might create a threshold which limits adjustment in the case of shocks that are too small to imply that adjustments are profitable. Threshold unit root techniques are especially relevant in situations where deviations from equilibrium depend on transaction costs. The approach is to extend the univariate threshold unit root testing procedure of Caner and Hansen (2001) to a panel data context by aggregating the individual results with Fisher type tests of Maddala and Wu (1999) and Choi (2001).

Assuming that time series ($s_{i0}, \ldots, s_{iT}$) on the cross section units $i = 1, 2, \ldots N$ are generated for each i by an autoregressive process, the Augmented Dickey Fuller (ADF) regression is written as

$$\Delta s_{i,t} = \alpha_i + \rho_i s_{i,t-1} + \sum_{j=1}^{k_i} \gamma_{ij} \Delta s_{i,t-k} + \varepsilon_{i,t} \tag{5}$$

where $s_{i,t}$ is the log of the rent-value ratio for each state, $\varepsilon_{i,t}$ is an *i.i.d.* error term, and k_i is the autoregressive lag order included to ensure that the regressions residuals behave like white-noise processes. [2] The null of unit root is rejected if ρ_i is statistically significant and less than zero.

[2] The procedure of lag order selection used in this paper is the minimization of the Akaike information criterion (AIC) defined as: $AIC = -2(l/T) + 2(k/T)$, where l is the log of the likelihood function with k parameters estimated using T observations.

Empirical researchers are faced with the fact that the conventional unit root tests are usually unable to reject the hypothesis that spread is nonstationary. This might be because the conventional methods cannot disentangle nonstationarity from nonlinearity because of the joint modeling problem of unit roots and thresholds. Caner and Hansen (2001) propose a procedure that allows testing for unit root against the stationary two-regime threshold specification. Differing from their univariate specification, the model here is specified with the cross-section subscript identifiers in order to extend the tests to a panel data context:

$$\Delta s_{i,t} = \theta_{i1}^{'} x_{i,t-1} 1_{\{Z_{i,t-1} < \lambda_i\}} + \theta_{i2}^{'} x_{i,t-1} 1_{\{Z_{i,t-1} \geq \lambda_i\}} + e_{i,t} \tag{6}$$

where *t=1,.....T*, *i=1,....,N*, $x_{i,t-1} = (s_{i,t-1} d_{i,t}^{'} \Delta s_{i,t-1} ....\Delta s_{i,t-k})^{'}$ is a vector of right hand side variables, $1_{\{\ \}}$ is the indicator function, $e_{i,t}$ is an i.i.d error, $Z_{i,t} = s_{i,t} - s_{i,t-m}$ for some delay parameter $m \geq 1$. In this paper, m is restricted to be either one or two because of the limited time series dimension of the data. Note that the threshold variable $Z_{i,t-1}$ satisfies the stationarity and ergodicity conditions described in section 2 of Caner and Hansen (2001). The threshold variable is defined in changes rather than levels since econometric theory developed in Caner and Hansen (2001) does not allow levels. $d_{i,t}^{'}$ is a vector of deterministic components including an intercept and possibly a linear time trend. Both cases of constant and constant and a trend are considered in this study. The threshold λ_i is unknown and must be estimated. It takes on values in the interval $\lambda_i \in \Lambda[\lambda_{i1}, \lambda_{i2}]$ where λ_{i1} and λ_{i2} are picked so that $P(Z_{i,t} \leq \lambda_{i1}) = .15$ and $P(Z_{i,t} \leq \lambda_{i2}) = .85$. The coefficient vectors in each regime can be decomposed as

$$\theta_{i1} = \begin{pmatrix} \rho_{i1} \\ \beta_{i1} \\ \gamma_{i1} \end{pmatrix} \quad \text{and } \theta_{i2} = \begin{pmatrix} \rho_{i2} \\ \beta_{i2} \\ \gamma_{i2} \end{pmatrix}$$

For each $\lambda_i \in \Lambda$, equation (6) is estimated via OLS to obtain an estimated residual variance for each possible threshold value. The least-squares estimate of the threshold ($\widehat{\lambda}_i$) is found by minimizing this residual variance, and the other parameters estimates are then found by plugging in this threshold value into equation (6).

The second stage of the approach is to test for the non-linearity hypothesis against the null of linearity. To this purpose Caner and Hansen (2001) propose a simple Wald test (W_T) of equation (6) against a linear alternative, which is found to have a nonstandard distribution under the null, partially due to the presence of a parameter that is not identified under the null (a nuisance parameter), and partially due to the assumption of a nonstationary autoregression. To approximate the sampling distribution the authors propose two bootstrap procedures, one is for the stationary case, and the other is for the unit root case. Since the true order of integration is unknown, the authors suggest calculating the bootstrap p-values for both cases, and base inference on the more conservative (larger) p-value.

The last stage of the Caner and Hansen (2001) procedure involves testing for unit roots. Here the null hypothesis is the existence of a unit root but, there are two different alternative hypotheses considered.

$$H_{i0}: \rho_{i1} = \rho_{i2} = 0$$

A major case of interest is if the spread follows a stationary, threshold, autoregressive pattern, so the alternative for this case is:

$$H_{i1}: \rho_{i1} < 0, \rho_{i2} < 0$$

Another case in between H_{i0} and H_{i1} is the case of partial unit root. The alternative of this case can be expressed as:

$$H_{i2}: \begin{Bmatrix} \rho_{i1} < 0 and \rho_{i2} = 0 \\ \rho_{i1} = 0 and \rho_{i2} < 0 \end{Bmatrix}$$

So if H_{i2} holds then log of rent-value is nonstationary but not in a classic unit root fashion (Caner and Hansen, 2001). The test statistics for testing H_{i0} versus H_{i1} is a simple one sided Wald statistics defined in section 5 of Caner and Hansen (2001) as:

$$R_{i1T} = t^2_{i1\{\widehat{\rho}_{i1}<0\}} + t^2_{i2\{\widehat{\rho}_{i2}<0\}}$$

where t_{i1} and t_{i2} are the t ratios for $\widehat{\rho_1}$ and $\widehat{\rho_2}$ respectively from the estimation of equation (6). To discriminate between the stationary case (H_{i1}) and the partial unit root cases (H_{i2}), the authors suggest examining the individual t-statistics. If only one of these t-statistics is statistically significant, the process is consistent with the partial unit root cases. To obtain the significance levels for these unit root tests Caner and Hansen (2001) derive asymptotic approximations to the distributions of the statistics. However, the authors suggest using bootstrapping to obtain improved finite sample inference. Bootstrap samples must be constructed under the null of a unit root.

In order to extend the Caner-Hansen approach to a panel data context, I propose the use of the Maddala and Wu (1999) and Choi (2001) Fisher type tests to combine the individual results of the Caner –Hansen approach into panel measures of their statistics. Let $pval_i$ be the exact p-value of one of the Caner and Hansen (2001) tests for state i. The panel test statistics considered are:

$$P^{MW} = -2\sum_{i=1}^{N} In(pval_i)$$

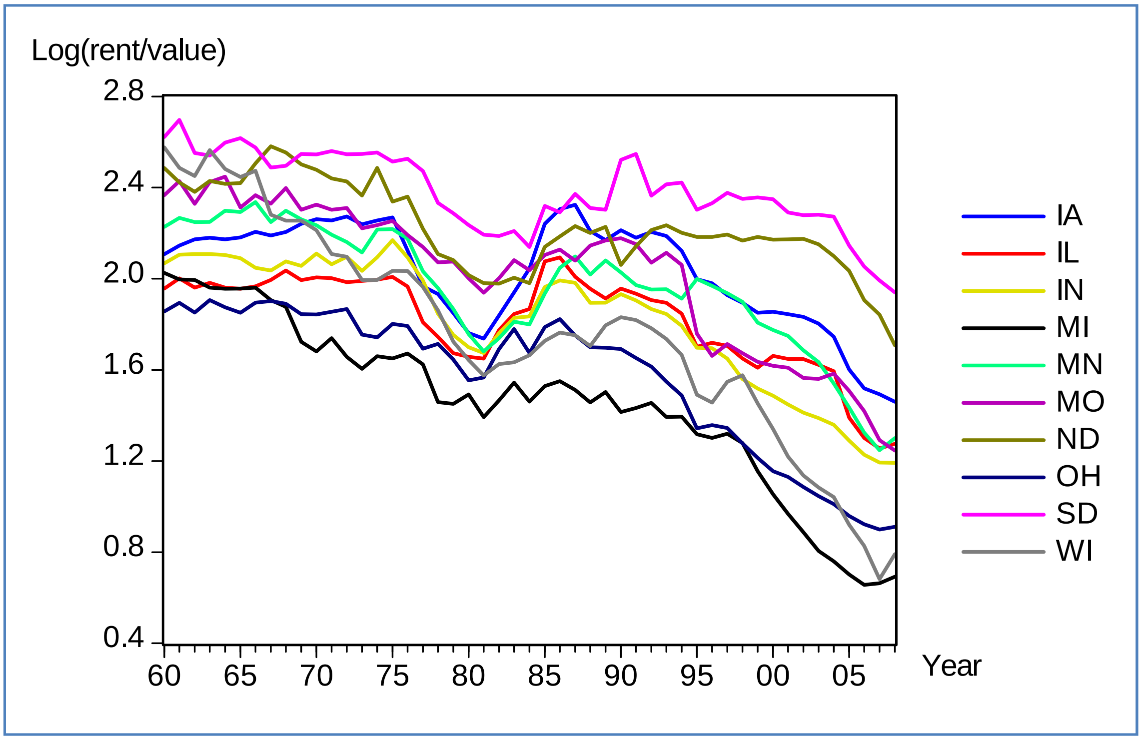
$$Z^{C} = \frac{1}{\sqrt{N}}\sum_{i=1}^{N} \Phi^{-1}(pval_i)$$

where $\Phi^{-1}(.)$ is the inverse of standard normal cumulative distribution function. P^{MW} is suggested by Maddala and Wu (1999) whereas Z^{C} is suggested by Choi (2001). The null hypothesis implies that the spread is nonstationary in all states. The alternative hypothesis states that spread is stationary for some of the states. These two tests are typical combination tests that are used often in meta-analysis. These panel unit root tests have several advantages over other tests. First, they do not need the assumption that all the series follow the same process, as the pooled unit root test of Levin, Lin and Chu (2002) does. They do not require a balanced panel, as in the case of the Im, Pesaran and Shin (2003), and one can use different lag lengths and exogenous variables in the individual regressions. Finally, the usefulness of these tests is not restricted to the ADF unit root test, as they can be derived from any univariate unit root test that leads to a continuous test statistics. There are two important restrictions of these panel tests: first, the p-values have to be derived by simulation (i.e., they have to be exact p-values), and second, the tests are valid only under the crucial assumption of cross-sectional independence. The first restriction is met as the p-values of Caner-Hansen tests are derived by simulations. The second point is an important topic of current research on panel unit root tests (Breitung and Pesaran, 2005). The issue of the cross sectional dependence is discussed in the next section; however, it is assumed throughout the paper that the states are cross sectionally independent.

4. Data and Empirical Findings

Annual observations for 10 agricultural states in the U.S between 1960 and 2008 are used for this study. The USDA classifies agricultural states under different regions. These regions are not all similar in their agricultural activities and land markets. Therefore, one constraint in selecting the states was to look at 'similar' states in order to reduce the complexities of working with a very heterogeneous panel data set. The availability of data for cash rents was another critical constraint in determining which states to include. The three regions included in the analysis are Lake States (Michigan, Minnesota and Wisconsin), Corn Belt (Illinois, Indiana, Iowa, Missouri and Ohio) and Northern Plains (North Dakota and South Dakota). Farmland prices are based on estimates of the value of land and buildings per acre, obtained from the US Department of Agriculture, National Agricultural Statistics Service and the Economic Research Service. Rents are based on cropland cash rents per acre and come from the same sources.[3] Both series are based on opinion surveys. The analysis is based on the log of the rent to value ratio. This series for each of the ten states can be followed in Figure 1. We can observe the boom and boost cycles in land values throughout the sample period as well as the negative trend in the spread suggesting a higher land price relative to cash rents from the land. The question is whether spread is stationary in the long run as suggested by the PVM.

Figure 1. Log of rent-to-value ratio (spread), 1960-2008



[3] The cash rents and price variables are not entirely consistent because of the availability issues with the data. Land values and cash rents are reported by ERS up to 1996 and by NASS afterwards. Even though the cash rents data for all farm real estate are available up to 1996, they are not available after that date. Since a long time dimension is needed for a robust analysis, I used cropland cash rents as the cash rents series. This decision can be justified by the fact that the ratio of the two cash rents series is stationary over their common period of 1967-1996. ERS's data set on value of land and buildings does not include the value of dwellings, but NASS's data does include dwellings. An adjustment has not been made, because the adjustment causes more than 5% difference in the common years' data. Nebraska and Kansas have been excluded even though they are in the Northern Plains. This is because the cash rents data are reported separately for irrigated and non-irrigated cropland in Kansas and Nebraska, and the rates of irrigation are higher in those states than in North and South Dakota. Further, irrigated cropland cash rents do not start until 1975 in Kansas and not until 1970 in Nebraska. The values of land and buildings are available from 1960, but cropland cash rents for the ten states start from 1967. The cropland cash rents series has been projected backwards to1960, using a linear trend model with a constant and using the farm real estate cash rents series as a benchmark since it is available from 1960.

Table 1. Conventional Univariate Unit Root Tests

	ADF						PP				KPSS (H0: no unit root)			
	Constant			Constant and Trend			Constant		Constant and Trend		Constant		Constant and Trend	
State	ρ	test stat (p-val)	lag	ρ	test stat (p-val)	lag	adj. test stat (p-val)	band width	adj. test stat (p-val)	band width	LM stat	band width	LM stat	band width
Lake States														
MI	0.001	0.027 (0.956)	4	-0.175	-2.222 (0.466)	3	0.292 (0.976)	4	-1.529 (0.806)	4	0.842 [0.463]	5	0.137 [0.146]*	5
MN	-0.013	-0.324 (0.913)	1	-0.174	-2.479 (0.337)	2	0.137 (0.965)	3	-1.580 (0.786)	3	0.719 [0.463]	5	0.092 [0.146]*	5
WI	-0.002	-0.068 (0.947)	1	-0.140	-1.905 (0.636)	1	0.099 (0.963)	0	-1.543 (0.800)	1	0.853 [0.463]	5	0.106 [0.146]*	5
Corn Belt														
IL	-0.031	-0.598 (0.861)	1	-0.132	-1.852 (0.663)	1	-0.026 (0.951)	2	-1.486 (0.821)	2	0.637 [0.463]	5	0.137 [0.146]*	5
IN	-0.002	-0.083 (0.945)	1	-0.111	-1.897 (0.640)	1	0.482 (0.984)	2	-1.697 (0.737)	3	0.779 [0.463]	5	0.148 [0.146]	5
IA	-0.050	-1.244 (0.647)	1	-0.111	-2.285 (0.433)	1	-0.513 (0.880)	4	-1.639 (0.762)	4	0.500 [0.463]	5	0.111 [0.146]*	5
MO	0.027	0.803 (0.993)	0	-0.117	-1.446 (0.834)	0	0.870 (0.994)	1	-1.402 (0.848)	2	0.823 [0.463]	5	0.150 [0.146]	5
OH	0.025	0.956 (0.996)	0	-0.082	-1.436 (0.837)	0	0.956 (0.996)	0	-1.452 (0.832)	1	0.790 [0.463]	5	0.197 [0.146]	5
N. Plains														
ND	-0.029	-0.486 (0.885)	2	-0.171	-2.060 (0.554)	2	-0.463 (0.889)	3	-1.626 (0.768)	3	0.632 [0.463]	5	0.094 [0.146]*	5
SD	-0.044	-0.667 (0.845)	0	-0.177	-1.830 (0.674)	0	-0.471 (0.888)	2	-1.902 (0.638)	3	0.651 [0.463]	5	0.085 [0.146]*	5

Bandwidth is selected via Barlett Kernel and NeweyWest. For both ADF and PP tests, the critical values of MacKinnon (1991) are used to test the unit root hypothesis. For the KPSS test, null hypothesis is stationarity. Numbers in brackets reported under KPSS LM-test statistics are 5% critical values.

In this section we use the empirical framework given in section 3 and try to understand whether this new econometric methodology will further our knowledge in understanding the behavior of farmland values. Before beginning the nonlinear tests, I consider conventional univariate tests for unit roots against linear stationary alternatives. The most commonly applied tests are Augmented Dickey Fuller (ADF) (Dickey and Fuller, 1979) and Phillip Perron (PP) (Phillips and Perron, 1988) tests. They differ in how they treat serial correlation in the test regressions. ADF tests adopt a parametric autoregressive structure to capture serial correlation while PP tests use non-parametric corrections based on estimates of the long-run variance of dependent variable.

Table 2. Linear Panel Unit Root Tests

	Exogenous variables: Individual effects		Exogenous variables: Individual effects, individual linear trends	
Method	**Test Statistic**	**P-value**	**Test Statistic**	**P-value**
Null: Unit root (assumes common unit root process)				
Levin, Lin & Chu t*	3.579	0.999	1.259	0.896
Breitung t-stat	-1.311	0.095*	-1.959	0.0250**
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	4.778	1.000	1.204	0.885
Maddala-Wu (ADF - Fisher Chi-square)	2.266	1.000	9.105	0.981
Maddala –Wu (PP - Fisher Chi-square)	1.077	1.000	5.031	0.999
Choi-(ADF-Fisher Standard Normal)	4.753	1.000	1.323	0.907
Choi-(PP-Fisher Standard Normal)	5.731	1.000	2.479	0.993
Null: No unit root (assumes common unit root process)				
Hadri Z-stat	13.019	0.000	4.500	0.000

Probabilities for Fisher tests are computed using an asympotic Chi-square distribution. All other tests assume asymptotic normality. * *, and * denote siginificance at 5% and 10% levels respectively.

KPSS test from Kwiatkowski et al. (1992) is especially useful in cases of smaller samples as the null hypothesis of this test is stationarity instead of a unit root. The results are reported in Table 1. ADF and PP tests can not reject the null hypothesis while KPSS accepts the null of stationarity for the 7 states out of 10 when a trend with a constant is included in the model. This might be a sign for the low power of ADF type unit root tests for our sample.

Panel unit roots tests with linear adjustment processes are reported in Table 2. For the most part, the results indicate the presence of a unit root. The Levin, Lin and Chu (2002) (LLC), Im, Pesaran and Shin (2003) (IPS), and both Fisher tests by Maddala and Wu (1999) and Choi(2001) fail to reject the null of a unit root. Similarly, the Hadri (2000) test statistic, which tests the null of no unit root, strongly rejects the null in favor of a unit root. The one exception to this pattern is the Breitung (2002), which does reject the unit root null.

Table 3 gives test results for a threshold in the spread in 10 U.S. states. Bootstrap p-values and the corresponding optimal delay parameters along with the observations that fall into each regime. There is a little evidence of thresholds in the sample based on the Wald test statistics. Only Michigan, Ohio and South Dakota seem to have significant nonlinearities in their log of rent-value ratios. However, the coefficient of adjustment changes dramatically for Iowa, Missouri, Ohio and North Dakota as well. The sample splitting in the states with insignificant results is very unbalanced, suggesting that the analysis might be suffering from a small sample bias. Therefore, I suspect that there are more cases of nonlinearities than the formal tests are suggesting in Table 3. Threshold estimate is changing from 1 percent to 7 percent in states.

Table 3. Nonlinearity Tests of Caner and Hansen (2001)

	Constant						Constant and Trend					
			Observations in		Coefficient of adjustment				Observations in		Coefficient of adjustment	
State	**Threshold estimate**	**m**	**Regime 1**	**Regime 2**	**ρ1**	**ρ2**	**Threshold estimate**	**m**	**Regime 1**	**Regime 2**	**ρ1**	**ρ2**
Lake States												
MI	-0.061	2	12	34	0.133	-0.021	-0.061	2	12	34	0.0008	-0.134
	[0.026] *				(0.068)	(0.026)	[0.036] *				(0.151)	(0.076)
MN	-0.077	1	9	37	-0.017	-0.083	-0.077	1	9	37	-0.102	-0.278
	[0.490]				(0.130)	(0.045)	[0.148]				(0.280)	(0.068)
WI	-0.035	1	24	22	-0.010	-0.094	-0.036	1	23	23	-0.214	-0.218
	[0.458]				(0.041)	(0.048)	[0.465]				(0.108)	(0.117)
Corn Belt												
IL	0.045	2	41	5	-0.014	0.366	0.045	2	41	5	-0.130	1.934
	[0.412]				(0.059)	(0.281)	[0.238]				(0.075)	(0.701)
IN	0.033	2	36	10	-0.011	-0.327	0.033	2	36	10	-0.095	-0.609
	[0.259]				(0.031)	(0.143)	[0.272]				(0.063)	(0.187)
IA	0.018	2	32	14	0.004	-0.515	-0.023	1	19	27	0.107	-0.297
	[0.179]				(0.047)	(0.145)	[0.213]				(0.079)	(0.070)
MO	0.054	2	38	8	0.023	-0.182	-0.048	1	16	30	-0.232	-0.019
	[0.328]				(0.038)	(0.196)	[0.265]				(0.137)	(0.109)
OH	-0.057	1	11	35	-0.046	0.032	-0.070	1	6	40	-1.07	-0.117
	[0.067] *				(0.104)	(0.028)	[0.073] *				(0.513)	(0.056)
N. Plains												
ND	0.026	2	35	11	0.035	-0.243	0.071	1	40	6	-0.135	-1.425
	[0.510]				(0.065)	(0.115)	[0.349]				(0.080)	(0.477)
SD	0.019	2	34	12	0.109	-0.415	0.007	2	30	16	0.092	-0.712
	[0.066] *				(0.077)	(0.132)	[0.058] **				(0.115)	(0.170)

The numbers in brackets, below the threshold estimates, are the bootstrap p-values of W_T statistics (Wald test statistics for the threshold effect). The numbers in parentheses are the standard errors of the coefficients of adjustment. p-values are calculated using the bootstrap method in Section 4.3 of Caner and Hansen (2001). For bootstrapping, 10,000 replications have been used. Threshold estimate is λ in the model. m represents the optimal delay parameter and maximum m is restricted to be 2. 10,000 bootstrap repetitions are used. "**" and "*" denote significance at 5% and 10% levels respectively.

Table 4. Univariate Threshold Unit Root Tests

State	Constant			Constant and Trend		
	R1T	t1	t2	R1T	t1	t2
Lake States						
MI	0.893	0.979	0.596	0.823	0.844	0.470
MN	0.612	0.739	0.300	0.088*	0.791	0.022**
WI	0.536	0.701	0.266	0.482	0.371	0.456
Corn Belt						
IL	0.959	0.711	0.956	0.855	0.486	0.999
IN	0.418	0.697	0.181	0.206	0.574	0.101*
IA	0.085*	0.804	0.024**	0.092*	0.962	0.023**
MO	0.870	0.862	0.561	0.837	0.430	0.851
OH	0.951	0.662	0.938	0.367	0.318	0.378
N. Plains						
ND	0.506	0.851	0.254	0.237	0.472	0.159
SD	0.151	0.946	0.057**	0.064*	0.924	0.021**

R1T ; t1; t2 are unit root tests described in section 3 , specifically t1; t2 are tests for H0 versus H2 (stationarity in regime 1 *or* 2) and R1T is for testing H0 versus H1 (stationarity in regime 1 *and* 2). Under their respective columns reported are the bootstrap p-values. 10,000 bootstrap replications are used. “**” and "*" denote significance at 5% and 10% levels respectively.

The unit root test results assuming nonlinear models are reported in Table 4. One-sided Wald test (R1T) and t1; t2 tests for unit roots are used to determine whether the two regimes are nonstationary or not. One-sided Wald tests, which test unit roots against a two-regime stationary nonlinear model, are rejected in Minnesota, Iowa and South Dakota. However, looking at the results of the individual t tests in each regime reveals that the significance of Wald statistics is due to the stationarity of the second regime. Similarly, the Wald test can not reject the unit root in both regimes in Indiana; however the trended case suggests stationarity in the second regime.

Table 5. Panel threshold unit root tests

	Exogenous variables: Individual effects		Exogenous variables: Individual effects, individual linear trends	
Method	**Test Statistic**	**P-value**	**Test Statistic**	**P-value**
Nonlinear Maddala-Wu (ADF - Fisher Chi-square)	14.736	0.791	25.696	0.176
Nonlinear Choi (ADF-Fisher Standard Normal)	1.119	0.868	-0.998	0.159

The last step of the empirical analysis is to combine the information obtained from univariate threshold procedures with Maddala and Wu (1999) and Choi (2001) statistics. Panel version of the threshold unit root tests reported in Table 5 do not show evidence in favor of the stationarity of log rent-value ratio in US farmland markets. However, note that the threshold version of the Fisher type test statistics reported in Table 5 are comparable to those in Table 2 (ADF counterparts). Comparing the test statistics and p-values suggests that even though incorporating nonlinearities into the model does not change the result that spread follows a nonstationary process, it certainly improves the significance of the test statistics. Especially, the trended cases in Table 5 are very close to rejecting the null of a unit root.

The empirical results found in this paper are far from validating the implications of Present Value Model discussed in Section 2; however they can give a direction for further research on the issue. There are two things that might cause poor empirical results when testing the stationarity of log (rent/value) ratio: cross section dependence, and small samples.

Cross section dependence can arise due to a variety of factors, such as omitted observed common factors, unobserved common factors, or general residual interdependence that could remain even when all the observed and unobserved common effects are taken into account. Pesaran (2004) proposes a simple test for error cross section dependence that has correct size and sufficient power even in small samples. To check if the panel at hand is characterized by cross section dependence, the residuals of the individual ADF regressions from the preceding single state analysis are used to compute Pesaran's (2004) test statistic. The test statistic of cross section dependence is computed as

$$CD = \sqrt{\frac{2}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\sqrt{T_{i,j}}\,\hat{\pi}_{i,j}\right)$$

where $\hat{\pi}_{i,j}$ are the pairwise correlation coefficients from the residuals of the ADF regressions. The CD test statistics reported in Table 6 clearly indicate the existence of cross section dependence. The hypothesis of zero cross section correlation is rejected in all cases except the

Northern Plains. One approach to overcome the cross section dependence is to modify the ADF type panel unit root regression by including the cross section means of the right hand side variables as additional explanatory variables (Pesaran, 2007). The other approach is using bootstrapping as suggested by Wu and Wu (2001).

Table 6. Test of Cross Section Dependence within Different Regions

	All states	**Lake States**	**Corn Belt**	**N. Plains**
Residuals from ADF regression with intercept				
CD-statistics	17.598	4.974	11.865	0.888
p-value	0.000	0.000	0.000	0.374
$\hat{\bar{\rho}}$	0.193	0.212	0.276	0.065
Residuals from ADF regression with intercept and linear trend				
CD-statistics	17.339	4.638	11.979	0.760
p-value	0.000	0.000	0.000	0.446
$\hat{\bar{\rho}}$	0.190	0.197	0.279	0.056

Small sample size is another problem, because the approach in this paper is based on combinations of univariate unit root tests and sample splitting (Caner and Hansen, 2001) which makes it important to have a long time span in the data. Since we can't extend the time series dimension of the data, we will adopt another approach in testing the panel stationarity. Testing for a common unit root and assuming a common threshold variable for all states may be a better way of dealing with the sample size issue.

Another possibility might be that a threshold variable in levels instead of the change in the log of rent-to-value might be more consistent with the data. A natural way to proceed in this case would be adopting a self exciting threshold autoregression (SETAR) instead of Caner and Hansen's (2001) threshold autoregression procedure. A future study aims to address all these issues together combining a fixed effects type of model with an appropriate nonlinear regime switching model.

5. Concluding Remarks

Because fluctuations in farmland prices can have serious consequences for the financial wellbeing of the sector, numerous studies have analyzed their behavior. However, earlier empirical studies usually fail to accept the hypotheses implicated by the Present Value Model. One possible explanation for the lack of consensus about farmland pricing and the explanatory power of PVM might be the presence of market frictions. Transaction costs and the large capital investments necessary to participate in the agricultural land market may cause nonlinearities in the adjustment towards long run equilibrium.

In this study, it is proposed to modify the conventional unit root tests of spread (log of rent-value ratio) to account for thresholds or other discontinuities in adjustment in order to increase the power of the tests in favor of rejecting the null of nonstationarity. A threshold autoregressive panel unit root approach is implemented extending the technique introduced by Caner and Hansen (2001) to a panel data context using Fisher-type tests as proposed by Maddala and Wu (1999) and Choi (2001). Empirical analysis using a panel of ten agricultural states of U.S. reveals little evidence on nonlinearities in the land markets. Only Minnesota, Iowa and South

Dakota show significant evidence in favor of a stationary threshold model against a unit root process in the log rent-to-value ratio. However, the results might be due to cross section dependence among the states or due to the fact that time span of the data is too short for a univariate, sample splitting based panel analysis. Future research will address these issues.

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