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# **THE PRESENT VALUE MODEL, FARMLAND PRICES AND STRUCTURAL BREAKS**

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# **THE PRESENT VALUE MODEL, FARMLAND PRICES AND STRUCTURAL BREAKS**

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## **Abstract**

We review the constant discount rate present value model of farmland prices using non-stationary panel data analysis. We use panel unit root and cointegration analysis to test if the present value model holds for a sample of 31 U.S. States covering the period 1960-2000. Preliminary results indicate that farmland prices and cash rents are non-stationary and non-cointegrated assuming a constant discount rate. The absence of cointegration may be due to the presence of a regime shift representing a time-varying discount rate. To accommodate this possibility, we introduce new panel cointegration tests that allow for unknown regime shifts in the cointegration relationship. The results suggest that the cointegration hypothesis cannot be rejected if there is a regime shift. Thus, while the present value model of farmland prices must be rejected when the discount rate is presumed constant, it cannot be rejected once we allow for regime shifts representing a time-varying discount rate.

Key words, Farmland prices, Present value model, Non-stationary panel data analysis, Regime shift.

JEL classification, C22, C23, G12, Q14.

## Introduction

There has been an enormous amount of research that has analysed the relationship between farmland prices and current and expected future returns on these assets (see, e.g. Phipps, 1984; Alston, 1986; Featherstone and Baker, 1987; Lloyd and Rayner, 1991; Falk, 1991; Falk and Lee, 1998; Lence and Miller, 1999; and more recently Moss and Schmitz, 2003).

Even with the rapid advances in technology in agriculture during the twentieth century, farmland remains a critical factor in production. The importance of farmland is underscored by its dominant position in the agricultural balance sheet over time. However, despite the importance of land to agricultural production and its dominance in the agricultural balance sheet, economic models explaining changes in land prices have only met with limited success. Although the literature suggests that Ricardian rents and the standard capitalization formula predict long-run variations in land prices, short-run deviations in the form of asset pricing bubbles cannot a priori be rejected. In addition, the economic fundamentals of farmland value may be changing. First, the linkage between farmland values and sector solvency may directly impact the economic viability of the farm sector. Second, the structure of agricultural production is changing. Third, farmland values are influenced not only by farm-related supply and demand factors but also by nonfarm demands for farmland. Fourth, understanding the dynamics of farmland valuation is critical in understanding the impacts of agricultural and trade policy, and in forecasting farmland values, (e.g. Pesaran and Timmerman, 2004).

In this paper we use time series data to test if the present value model of farmland pricing holds for a panel of 31 States, 1960-2000. The financial econometrics literature utilizes unit root and cointegration analysis. Therefore, we critically examine whether the economic theory of farmland pricing needs to be revised or if new panel econometric tests are needed.

The present value model of stock prices requires that, if the dividend of the stock price possesses a unit root, then so must the price of the stock itself. Moreover, prices and dividends should be cointegrated and, if the discount rate can be assumed to be constant, there should be a stable cointegration vector. Falk (1991) has applied this analysis to Iowa farmland prices. Although land prices and rent movements are highly correlated, Falk rejects the present value model. Clark, Fulton and Scott (1993) also found that the simple asset pricing model did not hold. Finally, Tegene and Kulcher (1993) also reject the present value model while Clark, Klein and Thompson (1993) found some support for this model.

These studies have examined whether the time series behaviour of economic variables can be characterized by a unit roots and cointegration. In general, the analysis has been carried out by using conventional time series methods. The main problem is that these unit root and cointegration tests can have poor power in realistically small samples. To overcome this problem, recent research has developed unit root and cointegration tests for panel data. By adding the information contained in the individual time series, panel data methods are able to generate more powerful tests. However, Gutierrez and Erickson (2004) have demonstrated that conventional panel cointegration tests are generally inappropriate in the presence of structural change since they assume that the cointegration vector is stable over time. This finding is particularly relevant when testing the present value model since the cointegration vector may be unstable as the discount rate is likely to be time-varying. To overcome the problem of a time-varying cointegration vector, the recent literature has proposed several panel cointegration tests that can be employed to accommodate structural breaks.

We apply panel unit root and cointegration tests to the constant discount rate version of the present value model for U.S. farmland prices. We employ a panel of 31 U.S. States for which time series data are available on farmland prices and cash rents over the period 1960-2000. We show that while conventional panel unit root and cointegration tests tend to support rejection of the constant discount rate version of the present value model, when considering regime changes representing a time-varying discount rate, the present value model cannot be rejected. In the following section, we briefly review the present value model and the panel unit root and cointegration methods adopted for this paper. Section 3 presents the empirical results and section 4 concludes.

## Testing the present value model for farmland prices

The basic framework for our analysis is a present value model. This model relates  $P_{it}$ , the real price per acre of farmland in State  $i$  at period  $t$ , to  $D_{it}$ , the real per acre rent in the same State during

period  $t$  to repayment to the landowner at period  $t + 1$ . In this notation, the real rate of return on an acre of land in State  $i$  from period  $t$  to  $t + 1$  may be defined as

$$R_{it+1} \equiv \frac{P_{it+1} + D_{it+1}}{P_{it}} - 1. \quad (1)$$

In this model we make the canonical assumption of a constant expected real return on an acre of land. Therefore, if  $E_t$  is the expectation operator conditional on the information available at time  $t$ , then this constant expected real return per acre assumption is expressed  $E_t R_{it+1} = R_i$ . This implies that the expected price of an acre of farmland in period  $t$  can be written as

$$P_{it} = (1 + R_i)^{-1} E_t (P_{it+1} + D_{it+1}). \quad (2)$$

Following e.g. Campbell and Shiller (1987), the above expression can be solved over  $K$  periods, which, by the law of iterated expectations, yields

$$P_{it} = \sum_{k=1}^K (1 + R_i)^{-k} E_t D_{it+k} + (1 + R_i)^K E_t P_{it+K}. \quad (3)$$

To preclude rational bubbles, it is necessary to assume that the farmland price  $P_{it}$  does not grow too fast relative to  $R_i$ . This implies that the second term in (3) vanishes as  $K \rightarrow \infty$ . Therefore, we obtain the following a unique solution for the expected price of an acre of farmland

$$P_{it} = \sum_{k=1}^{\infty} (1 + R_i)^{-k} E_t D_{it+k}. \quad (4)$$

This expression is very instructive. It suggests that if  $D_{it}$  possesses a unit root then so must  $P_{it}$ . Moreover, it suggests that  $D_{it}$  and  $P_{it}$  should be cointegrated with cointegrating parameter  $1/R_i$ . To show this, subtract  $D_{it}/R_i$  from both sides of (4), which yields

$$S_{it} \equiv P_{it} - R_i^{-1} D_{it} = \sum_{k=1}^{\infty} R_i (1 + R_i)^{-k} E_t \Delta D_{it+k}. \quad (5)$$

The term  $S_{it}$  equals the difference between the real price of farmland and the ratio of real per acre rent payment to the expected real return. This is usually referred to as the “spread in the financial literature. The expression in (5) suggests that if  $D_{it}$  is a unit root process, then  $\Delta D_{it+k}$  must be stationary for all  $k > 0$  and consequently  $S_{it}$  must be a stationary variable. It follows that  $D_{it}$  and  $P_{it}$  should be cointegrated with the cointegrating parameter  $1/R_i$ . This analysis suggests that there exist a straightforward way in which the constant expected return version of the present value model can be tested empirically. In fact, since the model posits that  $D_{it}$  and  $P_{it}$  are cointegrated, a natural first step in such a test would involve testing the variables for unit roots. If the variables are found to be nonstationary, a test of the constant expected return model can be constructed by first regressing  $P_{it}$  on  $D_{it}$  and then subjecting the residuals to a unit root test. If the null is rejected, the constant expected returns model of farmland prices might be said to hold.

The theory of a stable long-run relationship between farmland and rents is widely accepted and generally used for land appraisal purposes. Yet, oddly enough, it has been extremely difficult to verify empirically. In fact, most time series evidence tends to reject the hypothesized cointegrating relationship between farmland prices and rents (see, e.g. Falk, 1991; Clark, Fulton and Scott, 1993; Tegene and Kuchler, 1993). There are many explanations for this apparent lack of consensus. For example, de Fontnouvelle and Lence (2002) argue that the presence of market frictions drives a wedge between the price at which outsiders are willing to buy land and the price at which landowners are willing to sell it. To study the effects of such frictions, the authors introduce a new theoretical framework that allows for the presence of transaction costs. Featherstone and Baker (1987) suggest the presence of rational speculative bubbles in which prices deviate from their fundamentals associated with rents and interests rates.

In this paper we argue that the absence of popular empirical support for the present value model can be explained in part by the low power inherent in the conventionally applied time series methods. Therefore, we investigate this possibility before embarking on a major revision of the economic theory.

The problem facing researchers engaged in testing the present value model is that a unit root process can be easily mistaken for a process that is stationary around a deterministic trend. Therefore conventional time series unit root and cointegration tests may have insufficient power against the alternative hypothesis, especially considering the shortness of the time series usually available. To alleviate this problem, recent research has resorted to using panel data. This approach is advantageous not only because it augments the power of the time series method but also because it leads to tests that are normally distributed. This stands in sharp contrast to the conventional time series case, where test distributions usually involve complicated functions of Brownian motion.

There are, however, at least two major problems with this panel data approach that need to be appropriately addressed before any serious test of the present value model can be mounted. First, U.S. farmland prices typically display strong “boom/bust” cycles. If cycles are common across states, as seems to be the case, farmland prices in different states will tend to be correlated with each other. This is problematic because most existing tests for unit roots and cointegration in panel data assume independence, or at least zero correlation, among the cross-sectional units. If this assumption is violated, then these tests suffer from nuisance parameter dependencies, causing size distortions and deceptive inference. To address such cross-section correlation, we employ some of the most recent methods for testing for unit roots and cointegration in cross-sectionally dependent panels. Second, as pointed out by Falk, the expected real rate of return may not be constant but time-varying. In contrast to the first problem, which could be readily alleviated by using existing panel data methods, the second is potentially more problematic. This is so because the above test procedure is no longer valid since the postulated relationship between  $D_{it}$  and  $P_{it}$  is now non-linear. Specifically, the problem is that standard cointegration tests for panel data require that the cointegration vector be time-invariant. However tests based on this assumption are inappropriate when the cointegration vector is subject the structural change. Thus, since most panel data test statistics are constructed as simple averages of  $N$  time series statistics, the same kind of problems can be expected to materialize also in the panel data context. Indeed, Gutierrez and Erickson (2004) show by Monte Carlo simulations that the problems documented in the time series literature not only remains, but can even be exaggerated, in panel data.

The problem is that none of the existing tests for cointegration in panel data are able to accommodate such structural change. One reason for this is that allowing for structural change in existing tests represents a quite thorny undertaking. This is because the distribution of these tests with structural breaks will critically depend on nuisance parameters as indicated by the location of the breaks. Thus, it would be extremely difficult to control for the numerous possible combinations of heterogeneous breaks that might occur when using these tests in the panel data environment. Two studies that have analyzed this issue are those of Gutierrez and Erickson (2004) and Westerlund (2005). These will be described in the next section.

## Tests for cointegration in panel data with regime shifts

In this section, we review and extend upon the work of Gutierrez and Erickson (2004) and Westerlund (2005), which develops tests for cointegration in panel data that are able to allow for structural change. As we shall see, this will be key in testing the present value model of farmland prices. We begin with the study of Gutierrez and Erickson (2004), which generalizes the tests of Gregory and Hansen (1996) to panel data. The tests of Gutierrez and Erickson (2004) are constructed to test the null hypothesis of no cointegration versus the alternative hypothesis that there is at least one individual that is cointegrated with a single structural change. To this end, the authors propose several statistics that are appropriate under various types of changes. Since we want to allow for a time-varying expected real rate of return, we are interested in the alternative hypothesis of cointegration when all the parameters of the model are allowed to change at an unknown point in time. This type of shift is henceforth referred to as a regime shift and we will use  $0 < t_i < 1$  to denote the location of this shift for each individual. Now, with a single break, the regime shift model can be represented using the following dummy-variable regression

$$P_{it} = \mathbf{a}_{1i} + \mathbf{a}_{2i} \mathbf{j}_{itz}^c + D_{it} \mathbf{b}_{1i} + D_{it} \mathbf{b}_{2i} \mathbf{j}_{itz}^s + u_{it}, \quad (6)$$

where

$$\mathbf{j}_{itz}^c = \begin{cases} 0 & \text{if } t \leq [Tt_i] \\ 1 & \text{if } t > [Tt_i] \end{cases} \text{ and } \mathbf{j}_{itz}^s = \begin{cases} 0 & \text{if } t \leq [Tt_i] \\ 1 & \text{if } t > [Tt_i] \end{cases}.$$

The first step in the Gutierrez and Erickson (2004) test procedure involves computing the  $p$ -values of the individual  $ADF^*$ ,  $Z_{\hat{\alpha}}^*$  and  $Z_{\tau}^*$  statistics of Gregory and Hansen (1996). In the second step, the panel statistics are constructed by using the combination of  $p$ -values method suggested by Maddala and Wu (1999). Towards this end, if we let  $p_i$  denote the  $p$ -value obtained for individual  $i$  when using one of the Gregory and Hansen test statistics and let  $\Phi$  denote the standard normal cumulative distribution function, then we have the following panel test statistics

$$P = -\frac{1}{\sqrt{N}} \sum_{i=1}^N (\ln(p_i) + 1),$$

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

$$L = \frac{1}{\sqrt{p^2 N/3}} \sum_{i=1}^N \ln\left(\frac{p_i}{1-p_i}\right).$$

The  $P$  statistic is a modification of the inverse chi-square test of Fisher (1932), while the  $L$  statistic is a modification of the logit test. The  $Z$  statistic is usually referred to as an inverse normal test. Under the assumption of cross-sectional independence, each of the three statistics reaches a standard normal distribution under the null hypothesis as  $T \rightarrow \infty$  and  $N \rightarrow \infty$ . Under the alternative hypothesis that there is at least one individual that is cointegrated, then  $P$  diverges to positive infinity while  $L$  and  $Z$  diverges to negative infinity.

By contrast, the test of Westerlund (2005) is based on the LM test of Harris and Inder (1994) and thus takes cointegration as the null hypothesis. The alternative hypothesis is that there is at least one individual that is not cointegrated. This test is more general than the test of Gutierrez and Erickson (2005) since it allows for an unknown number of breaks in that can be located at different positions for different individuals. For notational convenience, let  $M$  denote the number of breaks for each individual located at dates  $T_{ij} = [Tt_{ij}]$ , where  $j = 0, \dots, M+1$ ,  $0 < t_{ij} < 1$ ,  $T_{i0} = 1$ ,  $T_{iM+1} = T$  and  $0 < M < J$  for some predetermined upper boundary  $J$ . Now, let  $LM_{ij}$  denote the Harris and Inder statistic obtained by regressing  $P_{it}$  on a constant and  $D_{it}$  using the subsample ranging from  $T_{ij}$  to  $T_{ij-1}$ . Then, the panel LM test of Westerlund (2005) can be written as

$$Z(M) = \sum_{i=1}^N \sum_{j=1}^{M+1} LM_{ij}$$

Following an appropriate standardization, the  $Z(M)$  statistic reaches a standard normal distribution as  $T \rightarrow \infty$  and then  $N \rightarrow \infty$  sequentially. To retrieve the location of each break for each possible value on  $M$ , Westerlund (2005) suggests using the method of Bai and Perron (1998, 2003), which estimates the break dates by minimizing the sum of squared residuals. Once the estimated break points have been obtained together with the associated sums of squared residuals for each allowable  $M$  for every individual, the unknown number of breaks can be estimated using an information criterion. Moreover, although we use  $M$  here to denote the number of breaks for all individuals, there is nothing that prevents  $M$  from varying across the panel. Thus, it is possible to have a situation in which  $M_i \neq M_j$  for  $i \neq j$ .

Westerlund (2005) implements the test for the case when the deterministic component of the cointegration regression is subject to structural change and provides Monte Carlo results to suggest that the test performs well in small samples. However, we are interested in the more general regime shift model in which not only the constant but also the individual slope parameters are permitted to vary.

Fortunately, extending the  $Z(M)$  statistic to the case with regime shifts is relatively straightforward and requires only modifying the break date estimator as suggested by Bai and Perron (1998, 2003).

To examine the small-sample properties of the  $Z(M)$  test in the regime shift model, we engage in a small Monte Carlo simulation exercise. The data generating process used for this purpose is that described in (6) with the error  $u_{it}$  being generated as the MA(1) process  $u_{it} = e_{it} + \mathbf{q}e_{it-1}$ , where  $e_{it} \sim N(0,1)$  and  $\mathbf{q} = (-0.5, 0, 0.5)$  is the moving average parameter. The parameters of the model are all generated as  $N(1,1)$  with the location of the structural break given by  $\mathbf{t}_i = (0.3, 0.5, 0.7)$ . The data are generated for 1,000 panels with  $N$  cross-sectional and  $T + 50$  time series observations. We disregard the first 50 observations for each cross-section in order to reduce the effect of the initial condition. In computing the fully modified residuals used for constructing the individual Harris and Inder (1994) test statistics, we use the Bartlett kernel with the bandwidth parameter set equal to  $[T^{1/3}]$ . The maximum number of breaks considered is  $J = 5$  and we use the Schwarz Bayesian information criterion for the estimation. The results are based on the asymptotic moments provided by McCoskey and Kao (1998).

<b>Table 1. Empirical size of the <math>Z(M)</math> test</b>						
<b>Unknown break point</b>						
$\hat{\delta}$	$T$	$N = 10$			$N = 20$	
		$\mathbf{q} = -0.5$	$\mathbf{q} = 0.0$	$\mathbf{q} = 0.5$	$\mathbf{q} = -0.5$	$\mathbf{q} = 0.0$
0.3	100	0.010	0.000	0.000	0.000	0.000
0.3	200	0.003	0.003	0.000	0.000	0.000
0.5	100	0.000	0.000	0.000	0.000	0.000
0.5	200	0.003	0.003	0.000	0.003	0.000
0.7	100	0.000	0.007	0.000	0.000	0.003
0.7	200	0.007	0.000	0.003	0.003	0.000
<b>Known break point</b>						
$\hat{\delta}$	$T$	$N = 10$			$N = 20$	
		$\mathbf{q} = -0.5$	$\mathbf{q} = 0.0$	$\mathbf{q} = 0.5$	$\mathbf{q} = -0.5$	$\mathbf{q} = 0.0$
0.3	100	0.007	0.020	0.000	0.003	0.057
0.3	200	0.003	0.020	0.000	0.017	0.023
0.5	100	0.000	0.020	0.000	0.003	0.033
0.5	200	0.010	0.020	0.000	0.013	0.023
0.7	100	0.013	0.017	0.000	0.010	0.020
0.7	200	0.007	0.030	0.000	0.010	0.027
Notes, The table reports the rejection frequencies on the 5% level. The value $\hat{\delta}$ refers to the location of the structural break. The value $\mathbf{q}$ refers to the moving average parameter. The results are based on 1,000 replications.						

The results on the size of a nominal 5% level test are presented in table 1. The results suggest that the test generally performs well but that it can be somewhat under-sized, especially in the case when  $\mathbf{q} = 0.5$ . Moreover, the effect of estimating the breaks rather than treating them as known leads to a reduction in the size and hence to a more conservative test. As expected from the asymptotic theory, we see that the performance under the null is unaffected by the location of the structural break. The results on the correct selection frequency suggest that the frequency count of the correctly chosen break locations when the estimated number of breaks is equal to its true value generally hovers around 60%. Based on this evidence, we conclude that it should be possible to successfully implement the  $Z(M)$  test even in the regime shift model.

## Empirical results



Our purpose is to provide some empirical evidence on the constant discount rate present value model for a panel comprised of 31 U.S. States sampled, 1960-2000, using the panel data methodology briefly presented in the previous sections. Thus, in this case,  $P_{it}$  will denote farmland prices measured as the estimated average value of an acre and  $D_{it}$  will denote cash rents per acre. Farmland prices are based on estimates of value of land and building per acre, U.S. Department of Agriculture, National Agricultural Statistics Service (NASS) and the Economic Research Service (ERS). Cash rents per acre are from the U.S. Department of Agriculture, NASS and ERS. All the series are deflated using the aggregate consumer price index (chain-weighted GNP deflator, 1996=100) (see Appendix 1).

We begin with the results on the unit root tests, which are presented in table 1. Three panel unit root tests are employed. They are the  $t$ -bar statistic of Im et al. (2003) and the  $t_a$  and  $t_b$  statistics of Moon and Perron (2004). All statistics are constructed under the null hypothesis that all of the cross-sectional units in panel are non-stationary against the alternative that there is at least one unit in the panel that is stationary. The difference lies in the assumptions made regarding the presence of cross-sectional correlation. The test of Im, Pesaran and Shin (2003) is the most restrictive since it presumes that the data are cross-sectionally independent. By contrast, Moon and Perron (2004) consider a linear dynamic factor model in which the panel is generated by both idiosyncratic shocks and unobserved dynamic factors that are common to all the units, thus explicitly permitting correlation among the cross-sectional units. Not accounting for cross-correlation in panel analysis can cause size distortions and deceptive inference on the unit root hypothesis. Hence, with strong dependencies on common cycles among States, accounting for such dependence may be important.

<b>Table 2. Panel unit root tests (a)</b>		
<b>Test</b>	<b>Farmland Prices</b>	<b>Cash Rents</b>
$t_a$	1.617(0.947)	1.945(0.974)
$t_b$	1.907(0.972)	2.511(0.993)
$t$ -bar	-3.974(0.000)	3.605(0.999)
(a) The numbers within parentheses are the $p$ -values.		

All statistics presented in the table are computed with an individual-specific constant in the test regression and we fix the order of the lag augmentation to two. Moreover, the values of the  $t_a$  and  $t_b$  statistics of Moon and Perron (2004) are calculated while permitting a maximum of three common factors. The appropriate number of factors was determined by using the  $BIC_3$  criterion developed by Bai and Ng (2002).

<b>Table 3. Panel cointegration tests (a)</b>	
<b>Test</b>	<b>Test value</b>
$ADF_p$	-2.507(0.994)
$ADF_z$	1.419(0.922)
$ADF_L$	1.282(0.900)
Panel $n$	0.744(0.228)
Panel $r$	2.061(0.980)
Panel $t$ (non-parametric)	2.829(0.998)
Panel $t$ (parametric)	9.945(1.000)
Group $r$	4.390(1.000)
Group $t$ (non-parametric)	5.272(1.000)
Group $t$ (parametric)	4.689(1.000)
LM	8.382(0.000)
(a) The numbers within parentheses are the $p$ -values.	

Save for the  $t$ -bar statistic, which leads to a rejection the null hypothesis of a unit root for farmland prices, the results suggest that we are unable to reject the null hypothesis. Hence, we conclude that both variables are non-stationary. To analyze if the rejection for farmland prices when using the  $t$ -bar statistic can be attributed to cross-sectional dependence, we compute the Pearson cross-correlation matrix of the residuals obtained from the individual augmented Dickey-Fuller regressions used in computing the statistic. The average cross-correlation is equal to 0.5 for farmland prices and 0.2 for cash rents. Thus, the rejection of the null for farmland prices may be attributed to a high degree of cross-correlation. We also computed the unit root tests proposed by Elliot, Rothemberg and Stock (1996) for each series of both variables. The results, which are not presented here, suggest that the null hypothesis of a unit root cannot be rejected.

Thus, since the variables appear to be non-stationary, we employ a battery of tests for cointegration. Table 3 presents the results for those statistics based on the assumptions of cross-sectional independence and no structural break. In particular, the table presents the results for the Panel and Group  $t$ ,  $r$  and  $n$  test statistics of Pedroni (1999, 2004), which are all based on the null hypothesis of no cointegration. The Panel statistics are designed to test this null hypothesis against the alternative hypothesis that all the individuals of the panel are cointegrated, whereas the Group statistics tests the null against the alternative that there is at least one individual that is cointegrated. As seen from the table, we are unable to reject the null hypothesis. This suggests that the constant present value model for farmland prices must be rejected. The table also presents the results obtained when using the Maddala and Wu (1999) combination of  $p$ -values method based on the parametric Group  $t$  statistic. These are denoted as  $ADF_P$ ,  $ADF_Z$  and  $ADF_L$ . As can be seen from the table, the results corroborate those obtained using the Pedroni (1999, 2004) tests.

All tests employed so far have been used to infer the null of no cointegration. However, since it is cointegration that is of primary interest, it is more natural to consider a residual-based test that takes cointegration as the null hypothesis rather than the opposite. Therefore, table 3 also presents some confirmatory results using the panel LM test proposed by McCoskey and Kao (1998). The test is based on the FMOLS estimator with the bandwidth parameter set equal to  $[T^{1/3}]$ . The results suggest that the null hypothesis must be rejected on all conventional levels of significance. Thus, all the test results contained in table 3 suggest that the present value model for farmland prices should be rejected.

<b>Table 4. Panel cointegration tests with regime shift (a)</b>					
<b>Test</b>	<b>Test value</b>	<b>Bootstrap distribution (b)</b>			
		1%	2.5%	5%	10%
$ADF_P^*$	5.97(0.00)	1.528	1.093	0.716	0.317
$ADF_Z^*$	-5.16(0.00)	-1.304	-1.074	-0.783	-0.503
$ADF_L^*$	-5.11(0.00)	-1.428	-1.148	-0.871	-0.590
$Z_{tP}^*$	6.02(0.00)	2.530	2.008	1.506	1.033
$Z_{tZ}^*$	-5.04(0.00)	-1.843	-1.391	-1.114	-0.777
$Z_{tL}^*$	-5.04(0.00)	-2.044	-1.596	-1.306	-0.961
$Z_{dP}^*$	6.05(0.00)	2.059	1.413	1.126	0.751
$Z_{dZ}^*$	-5.18(0.00)	-1.750	-1.470	-1.178	-0.820
$Z_{dL}^*$	-5.16(0.00)	-1.834	-1.545	-1.261	-0.925
$Z(M)$	1.89(0.03)	5.469	5.256	5.017	4.836
(a) The numbers within parentheses are the $p$ -values.					
(b) The bootstrap is based on 2,000 replications.					

Next, to analyze if this conclusion may be attributed to the presence of structural change or possibly cross-sectional dependence, we employ the tests proposed by Gutierrez and Erickson (2004), and Westerlund (2005). The results of the tests are reported in table 4 together with their bootstrapped distributions to allow for valid inference in the presence of cross-sectional dependence. For this purpose, we employ the sieve method proposed by Chang, Park and Song (2002). Also, the  $Z(M)$  test is

implemented using a maximum of three breaks with the appropriate number determined using the Schwarz Bayesian information criterion.

In contrast to the evidence based on the constant expected return version of the model, the results in table 4 suggest that the present value model cannot be rejected while entertaining the possibility of regime shift in the cointegration vector. Hence, if one allows for a time-varying expected real rate of return, then the present value model holds.

## **Conclusions**

In the paper we review the constant discount rate present value model of farmland prices using novel analysis of non-stationary panel data. As is well known, starting from the work of Campbell and Shiller (1987), the present value model can be formally tested using cointegration analysis. Specifically, the test requires first that farmland prices and cash rents are non-stationary variables and second that they are cointegrated with a stable cointegration vector. In recent years, both of these implications of the present value model have been scrutinized extensively using conventional time series unit root and cointegration methods. However, these tests are known to have deficient power, especially in realistically short samples. To remedy this problem, the recent literature on unit roots and cointegration has resorted to panel data. In this paper, we use panel unit root and cointegration analysis to test if the present value model holds for a sample of 31 U.S. States during the period 1960-2000. Preliminary results indicate that farmland prices and cash rents are non-stationary and non-cointegrated when the cointegrating vector is presumed to be time-invariant. Thus, we are able to reject the constant discount rate version of the present value model of farmland prices. We then demonstrate that this absence of cointegration may be attributed to the presence of one or more regime shifts representing a time-varying discount rate. Thus, while the present value model of farmland prices must be rejected when the discount rate is presumed constant, it cannot be rejected once we allow for a time-varying rate.

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## **Appendix 1. Data Sources**

### **States included in the study,**

*Northeast*, Delaware, Maryland, New Jersey, New York, Pennsylvania

*Lake States and Corn Belt*, Illinois, Indiana, Iowa, Michigan, Minnesota, Missouri, Ohio, Wisconsin

*Northern and Southern Plains*, Kansas, Nebraska, North Dakota, Oklahoma, South Dakota, Texas

*Appalachia*, Kentucky, North Carolina, Tennessee, Virginia, West Virginia

*Southeast*, Alabama, Florida, Georgia, South Carolina

*Delta*, Arkansas, Louisiana, Mississippi

### **Cash rents/acre,**

U.S. Department of Agriculture, NASS and ERS. *Agricultural Land Values and Cash Rents, Final Estimates*, USDA, NASS, Statistical Bulletins and ERS for earlier year's estimates.

### **Land and buildings values/acre,**

Estimates are based on estimates of land and building values (NASS) but ERS's estimates exclude the value of the operator's dwelling. U.S. Department of Agriculture, NASS and ERS. *Agricultural Land Values and Cash Rents, Final Estimates*, USDA, NASS, Statistical Bulletins and ERS for earlier year's estimates.

### **Gross Domestic Product (GDP) deflator,**

U.S. Department of Commerce, Bureau of Economic Analysis, 1996=100. Chain-weighted GDP deflator.