STRUCTURAL CHANGE, TRANSACTIONS COSTS, AND THE PRESENT VALUE MODEL OF FARMLAND: IOWA, 1900-1994

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Abstract

The present study makes the following contributions. First, two bootstrap tests of the constant-discount-rate present-value-model (CDR-PVM) in the presence of transactions costs are introduced. Second, such tests are applied to the longest series available on U.S. farmland prices and rents. Third, it is formally tested whether farmland returns experienced structural changes in this century. Farmland price behavior is found to be inconsistent with the CDR-PVM, even after considering typical transaction costs. The analysis also reveals that such inconsistency cannot be attributed to structural change. Therefore, the CDR-PVM failure must be explained by reasons other than transaction costs or structural change.

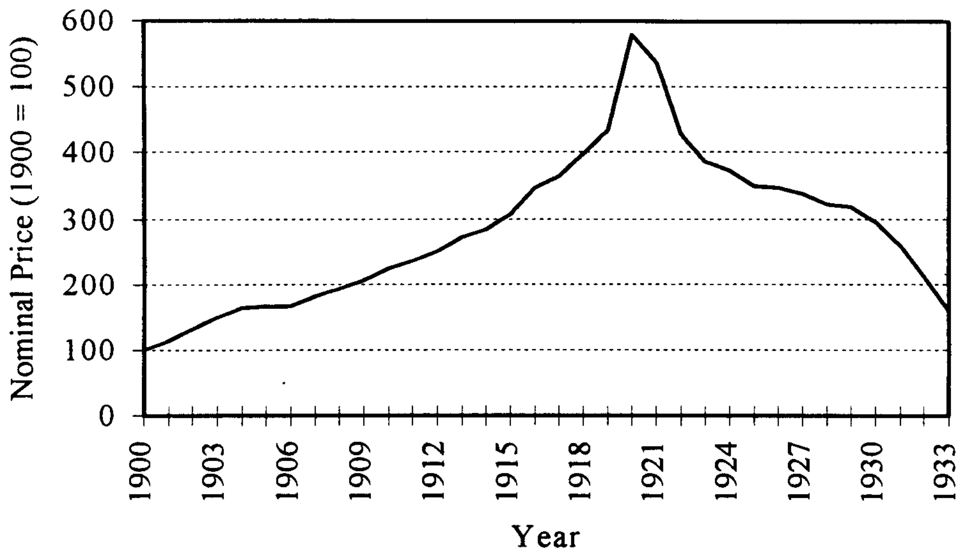
Introduction

Farmland prices have experienced notorious boom-bust cycles. For example, panels a and b of Figure 1 display the nominal prices of Iowa farmland over the two major boom-bust cycles that occurred in this century. The graph in panel a shows the price behavior over the first cycle. Land prices rose in every year from 1900 through 1920, for a total cumulative increase of 480% during the boom period. Subsequently, land prices fell in every year from 1920 through 1933; in 1933, prices were only about one-fourth of the price level achieved in the price peak of 1920. In the second cycle (panel b), land prices grew for 20 consecutive years from 1961 through 1981, for a total cumulative gain of 720%. After peaking in 1981, prices fell by almost two thirds in just 5 years.

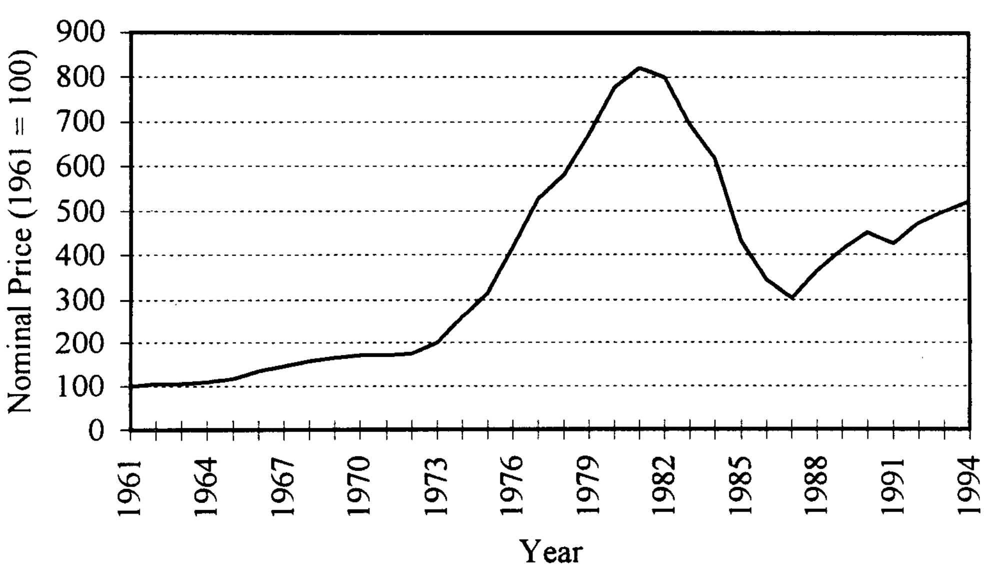
Farmland is the single most important asset in the farm sector, accounting for approximately 60% to 80% of the latter's aggregate balance sheet. Hence, the aforementioned boom-bust cycles in farmland prices have triggered noticeable wealth changes for the farm sector as a whole, as well as large wealth redistributions within it (Schmitz). Because land is also the most important source of collateral for agricultural lending, the availability of credit to the farm sector has been sharply curtailed at times of

Figure 1. Iowa farmland prices, 1900/33 and 1961/94.

a. Iowa farmland prices, 1900/33 (1900 = 100).



b. Iowa farmland prices, 1961/94 (1961 = 100).



decreasing land prices (Stam). As a result, the bust phase of the price cycle has been accompanied by great economic stress to the farm sector, as well as to the agricultural lending sector and to the rural economies in general.

Because of the importance of farmland values, numerous studies have analyzed their behavior. In the interest of space, such studies are not reviewed here because they are too numerous and have been surveyed elsewhere. Pope et al. present a highly critical assessment of the land pricing models most widely accepted prior to 1979. Robison and Koenig review developments in farmland price modeling up to 1990. The most recent literature is discussed at length by Weersink et al.

In general, most studies of farmland prices have pursued either a structural approach or a present-value-model (PVM) approach. The structural approach fell out of favor after Pope et al. showed that the corresponding empirical models could not explain the out-of-sample divergence between land prices and land rents. As a consequence, most of the current literature relies on the PVM, either directly or through the capitalization formula. The most important finding from the present-value studies is that the constant-discount rate (CDR) PVM of farmland pricing is consistently rejected (Falk; Clark, Fulton, and Scott; Tegene and Kuchler). Hence, the most widely used land pricing model seems inadequate to explain land price behavior.

Recently, Just and Miranowski developed a structural model that collapses to the CDR-PVM under numerous restrictions. Their model performs well when applied to state-level data from 1963 through 1986. However, the rational-expectations version of the model is outperformed by a version assuming naive expectations, thus reinforcing the results of present-value studies that reject the rational expectations hypothesis. Therefore, the efficient market hypothesis for farmland appears to be systematically violated. This is puzzling because it implies that excess returns could be achieved by trading in farmland based on knowledge of publicly available prices and rents.

Interestingly, the farmland asset pricing literature seems to have neglected the role of market frictions in shaping land price behavior. Such neglect is striking when considering the relatively large transaction costs involved in the transfer of farmland ownership (USDA 1964; Moyer and Daugherty; Thompson and Whiteside; Wunderlich).

Asset market frictions is an area of active research in financial economics (He and Modest; Luttmer; Aiyagari). An important theoretical finding from this research area is that market frictions may help explain some well-known asset-pricing puzzles.

As well, previous studies either have assumed no structural changes and used the longest data series available (e.g., Hanson and Myers), or have assumed the existence of a structural change at some arbitrary point and restricted analysis to subsequent data only (e.g., Shalit and Schmitz). However, none of the previous studies of farmland prices have conducted a formal test for structural changes in farmland returns. The most likely reason for this state of affairs is that there are several candidate points of change and appropriate search-based tests were recently developed (Andrews). Such information about structural changes in farmland returns should greatly improve our understanding of farmland prices.

Given the results from the existing literature, it is not clear whether the failure of the CDR-PVM is due to some fundamental flaw in the theoretical model, or simply due to the fact that the institutional setting of farmland markets (e.g., high transaction costs) is far removed from the theoretical assumptions underlying the CDR-PVM. Sorting between these two possible reasons for the CDR-PVM failure is important, as it would tell us in which direction we should proceed in our research agenda to gain a better knowledge about farmland price behavior.

We make the following contributions to the farmland pricing literature. First, we introduce two bootstrap tests of the theoretical implications of the CDR-PVM in the presence of transactions costs. Second, we apply such tests to the longest homogenous series available on U.S. farmland prices and rents. Third, we test whether farmland returns have experienced structural changes in the present century.

The study proceeds as follows. In the next section we present the theoretical model. This section is followed by a description of the econometric techniques. Next, the data used for the econometric tests are described. Empirical results are presented and discussed thereafter. The paper concludes with some final remarks.

Transaction Costs: Implications for Asset Pricing

In the absence of frictions, finance theory posits that the price of an asset is given by:

(3.1) $P_t = E_t[\delta_{t+1} (P_{t+1} + D_{t+1})]$,

where P_t is the real ex-dividend price of the asset at time t, δ_{t+1} is the stochastic discount factor,[1] $E_t(\cdot)$ is the expectation operator conditional on information at time t, and D_{t+1} is the real dividend paid by the asset at time t+1. According to (3.1), in market equilibrium the current price of an asset is equal to the expectation of the next period's asset payoffs (i.e., price plus dividends), where the latter are discounted by a stochastic factor.

Expression (3.1) is one of the most important results in finance theory, and is generally known as the "fundamental equation of asset pricing" (e.g., Dybvig and Ross) because it nests all of the standard asset pricing models. For example, Sharpe's CAPM, Breeden's Consumption CAPM, and Ross' APT can all be obtained from (3.1) by suitable specifications of the stochastic discount factor δ_{t+1} (Dybvig and Ross, Ingersoll). As well, (3.1) yields the CDR-PVM typically used in the farmland pricing literature (Falk):

(3.2) $P_t = \sum_{s=1}^{\infty} \delta^s E_t(D_{t+s})$.

Expression (3.2) is the stable forward solution of (3.1), obtained by recursive application of the latter after setting $\delta_{t+1} = \delta$, a constant.

Expression (3.1) is a necessary condition for asset market equilibrium. If the current asset price were smaller (greater) than the right-hand side of (3.1), agents would find it attractive to buy (sell) the asset now because doing so would yield an expected return above the required return (after accounting for risk); therefore, such a situation is inconsistent with market equilibrium.

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[1] $\delta_{i,t+1}$ is also called state-price density, pricing operator, pricing kernel, or intertemporal marginal rate of substitution. For von Neumann-Morgenstern preferences, $\delta_{i,t+1}$ is the ratio of the marginal utility of consumption for consecutive dates.

Asset Pricing under Transaction Costs: One-Period Holding Horizon

Although typically overlooked, an important assumption implicit in (3.1) is that the asset market is characterized by zero transaction costs. The expression analogous to (3.1) that must hold in equilibrium when there are transaction costs and agents plan to hold the asset for one period is derived intuitively next.[2] Let T_Π and T_Σ be the transaction cost that must be paid upon purchase and sale of the asset, respectively, expressed as a percentage of the asset price. In the presence of such costs, agents will buy (sell) the asset as long as expression (3.3) (expression (3.4)) holds:

$$(3.3) \quad (1 + T_\Pi)\, P_t < E_t\{\delta_{t+1}\, [(1 - T_\Sigma)\, P_{t+1} + D_{t+1}]\},$$

$$(3.4) \quad (1 - T_\Sigma)\, P_t > E_t\{\delta_{t+1}\, [(1 + T_\Pi)\, P_{t+1} + D_{t+1}]\}.$$

Therefore, in equilibrium it must be true that (3.5) and (3.6) hold simultaneously:

$$(3.5) \quad (1 + T_\Pi)\, P_t \geq E_t\{\delta_{t+1}\, [(1 - T_\Sigma)\, P_{t+1} + D_{t+1}]\},$$

$$(3.6) \quad (1 - T_\Sigma)\, P_t \leq E_t\{\delta_{t+1}\, [(1 + T_\Pi)\, P_{t+1} + D_{t+1}]\}.$$

Using the fact that $E_t[\delta_{t+1} D_{t+1}/(1 - T_\Sigma)] \geq E_t(\delta_{t+1} D_{t+1}) \geq E_t[\delta_{t+1} D_{t+1}/(1 + T_\Pi)]$, rearrangement of (3.5) and (3.6) yields the following expression that must hold in equilibrium:

$$(3.7) \quad \lambda^L \equiv \frac{-T_\Pi - T_\Sigma}{1 + T_\Pi} \leq E_t(h_{t+1}) \leq \frac{T_\Pi + T_\Sigma}{1 - T_\Sigma} \equiv \lambda^U,$$

where $h_{t+1} \equiv \delta_{t+1}\, (P_{t+1} + D_{t+1})/P_t - 1$.

Variable h_{t+1} is a stochastic discounted excess return. Albeit $h_{t+1} \neq 0$ in general, its conditional expectation $E_t(h_{t+1})$ must satisfy (3.7) for the market to be in equilibrium. By

[2]He and Modest provide a formal derivation for a specific class of asset pricing models.

comparing (3.1) and (3.7), it is straightforward to conclude that the former expression is a special case of the latter. If there are no transaction costs ($T_\Pi = T_\Sigma = 0$), $\lambda^L = \lambda^U = 0$ and expression (3.7) collapses to (3.1). In the presence of transaction costs (T_Π, $T_\Sigma > 0$), however, (3.7) indicates that there is a band of inaction given by $[\lambda^L, \lambda^U]$, inside which agents do not react to new information. If agents did react to new information within the band $[\lambda^L, \lambda^U]$, the expected gains from doing so would be more than offset by the losses stemming from the transaction costs.

Asset Pricing under Transaction Costs: Infinite Holding Horizon

As noted by Campbell, Lo, and MacKinlay (p. 316), however, agents can typically buy an asset and hold it for more than one period. In this instance, trading the asset may be profitable even if (3.5) and (3.6) hold simultaneously. As the planned holding horizon tends to infinity, the buying and selling conditions analogous to (3.3) and (3.4) become (3.8) and (3.9), respectively:

$$(3.8) \quad (1 + T_\Pi)\, P_t < \sum_{s=1}^{\infty} E_t(D_{t+s} \prod_{n=1}^{s} \delta_{t+n}),$$

$$(3.9) \quad (1 - T_\Sigma)\, P_t > \sum_{s=1}^{\infty} E_t(D_{t+s} \prod_{n=1}^{s} \delta_{t+n}).$$

Therefore, the testable condition becomes (3.10) rather than (3.7):

$$(3.10) \quad - T_\Sigma \leq \frac{1}{P_t} \sum_{s=1}^{\infty} E_t(D_{t+s} \prod_{n=1}^{s} \delta_{t+n}) - 1 \leq T_\Pi.$$

The CDR-PVM typically examined in farmland price studies (i.e., (3.2)) is a special case of (3.10). Expression (3.2) is obtained from (3.10) by letting transaction costs be zero ($T_\Pi = T_\Sigma = 0$) and by setting $\delta_{t+n} = \delta \;\forall\; n > 0$. In the absence of transaction costs, (3.10) says that the asset's sum of expected discounted dividends must equal the asset's current price. In the presence of transaction costs, the asset's sum of expected discounted dividends may be different from the asset's current price, but such difference should not exceed the transaction costs.

The infinite-holding-horizon model places more stringent restrictions on asset prices than the one-period holding horizon model. This can be demonstrated by using (3.5) and (3.6) recursively to obtain an expression analogous to (3.10), which yields much wider bounds on the pricing inequality. Given the very conservative nature of the one-period holding horizon model and the fact that land is typically held for many years (Rogers and Wunderlich), results from the infinite holding horizon model seem the most relevant for farmland.

Estimation Methods

Testing the CDR-PVM with One-Period Holding Horizon

The empirical analysis of the CDR-PVM assuming a one-period holding horizon is based on (3.7) with $h_{t+1} = \delta (P_{t+1} + D_{t+1})/P_t - 1$ (i.e., $\delta_{t+1} = \delta \ \forall\ t$). Note that if the series h_{t+1} contains either deterministic or stochastic trends, then the frictionless CDR-PVM cannot hold because $E_t(h_{t+1}) \neq 0$ in general, which violates (3.7) for $\lambda^L = \lambda^U = 0$. Therefore, the first step in testing the frictionless CDR-PVM consists of testing for the presence of deterministic and stochastic trends in the h_{t+1} series. If such trends exist, then (3.7) with $\lambda^L = \lambda^U = 0$ must be rejected.

If no trends are found in the h_{t+1} series, consider next estimating the autoregressive model of order p (AR(p)):

(4.1) $$h_{t+1} = \phi_0 + \phi_1 h_t + \ldots + \phi_p h_{t-p} + e_{t+1},$$

where the ϕs denote fixed coefficients and e_{t+1} is an i.i.d. error term. To be consistent with (3.7) in the presence of frictionless markets (i.e., $\lambda^L = E_t(h_{t+1}) = \lambda^U = 0$), estimation of (4.1) must yield $\phi_0 = \phi_1 = \ldots = \phi_p = 0$.[3] Hence, to test whether the frictionless CDR-PVM

[3]This assertion is true because $E_t(h_{t+1}) = \phi_0 + \phi_1 h_t + \ldots + \phi_p h_{t+1-p}$, which is generally different from zero unless $\phi_0 = \phi_1 = \ldots = \phi_p = 0$.

holds, here we propose estimating (4.1) and testing the null hypothesis H_0: $\phi_0 = \phi_1 = \ldots = \phi_p = 0$ against the alternative that the null is not true. Rejection of the null hypothesis provides strong evidence against the frictionless CDR-PVM. The proposed test draws on the one advocated by Enders (1988) to test for purchasing power parity.

Rejection of the null hypothesis H_0: $\phi_0 = \phi_1 = \ldots = \phi_p = 0$ does not imply rejection of the fundamental asset pricing model (3.7) in the presence of transaction costs ($\lambda^L < 0 < \lambda^U$). This assertion is true because the CDR-PVM in the presence of transaction costs can only be rejected if some of the calculated $E_t(h_{t+1})$s are significantly smaller than λ^L or significantly greater than λ^U.

Conceptually, a hypothesis test of the CDR-PVM in the presence of transaction costs is a multiple or induced hypothesis test (Savin). To see this, note that the implied null hypothesis may be stated as

$$(4.2)\quad H_0\text{: } \lambda^L \le E_p(h_{p+1}) \le \lambda^U \text{ and } \lambda^L \le E_{p+1}(h_{p+2}) \le \lambda^U \text{ and } \ldots \text{ and } \lambda^L \le E_N(h_{N+1}) \le \lambda^U,$$

where N is the number of observations in the sample. To find the appropriate rejection region for the test, we can use the complement of the $(1 - \alpha)\%$ joint confidence region for estimators of the conditional means. Let $\overline{H}_t^{\beta(\alpha)}$ ($\underline{H}_t^{\beta(\alpha)}$) be a sample statistic with a probability $[1 - \beta(\alpha)/2]\%$ of being greater (smaller) than the conditional expectation $E_t(h_{t+1})$. That is, for $p \le t \le N$, $[\underline{H}_t^{\beta(\alpha)}, \overline{H}_t^{\beta(\alpha)}]$ is a $[1 - \beta(\alpha)]\%$ confidence interval for $E_t(h_{t+1})$. Then, we can reject the CDR-PVM in the presence of transaction costs at the $\alpha\%$ significance level if there is at least one of the $(N - p + 1)$ confidence intervals (4.3) lying entirely below λ^L or above λ^U. This happens if and only if conditions (4.3) or (4.4) hold:

$$(4.3)\quad \lambda^L > \min[\overline{H}_p^{\beta(\alpha)}, \ldots, \overline{H}_N^{\beta(\alpha)}],$$

$$(4.4)\quad \lambda^U < \max[\underline{H}_p^{\beta(\alpha)}, \ldots, \underline{H}_N^{\beta(\alpha)}].$$

Although conceptually simple, actual implementation of the above test is far from trivial for a number of reasons. First, there is typically no closed-form solution for the $\beta(\alpha)$% significance level of the individual confidence intervals to make them consistent with an exact overall α% significance level for the test. To place an upper bound on the overall size of the test, we use the Sidák bound $\beta(\alpha) = 1 - (1 - \alpha)^{1/(N-p+1)}$ to approximate the individual significance levels. Although the resulting rejection regions may be somewhat smaller than required for an exact test, Monte Carlo studies indicate that the Sidák bounds are reasonably accurate and provide sharper inferences than well-known alternatives such as the Bonferroni bounds.

Further complications arise because specification tests typically provide limited support for the normality of the underlying noise distribution, and confidence bounds or rejection regions based on normal approximations may not reflect the finite sample properties of the data. As well, the least squares estimators of the AR(p) coefficients are known to be biased in finite samples, which further distorts the size of the multiple hypothesis tests. Therefore, we advocate using the bootstrap method proposed by Efron to solve these problems. Succinctly, our application of the bootstrap method to the hypothesis testing procedure involves the following steps:

1. Fit the AR(p) models for the h series.
2. Center the resulting residual vectors to form an empirical noise distribution $\hat{e}$.
3. For each of M bootstrap trials, draw (N – p) disturbances e^* from $\hat{e}$ (with replacement) to form bootstrap replicates of the data $h^*_{t+1} = \hat{\phi}_0 + \hat{\phi}_1\ h^*_t + \ldots + \hat{\phi}_p\ h^*_{t+1-p} + e^*_{t+1}$. Observed values of $h_1, \ldots, h_p$ are used to initialize the AR(p) sequence.
4. Compute the model parameters for each bootstrap sample $\phi^*_0, \ldots, \phi^*_p$.
5. Use bootstrap parameter estimates as an empirical distribution for $\hat{\phi}_0, \ldots, \hat{\phi}_p$, which are then used to approximate the distributions of the conditional means $E_t(h_{t+1}) = \hat{h}_{t+1}$.

Given the empirical distribution of conditional means for each time period, we can refer to the appropriate quantiles to form the joint rejection region with overall size α. To account for the bias of the conditional mean estimators, the quantiles are adjusted according to the BC procedure outlined in Section 4.1.3 by Shao and Tu. Let $H_{B,t}$ be the bootstrap cumulative distribution function (cdf) for $\hat{h}_{t+1}$ and Φ be the standard normal cdf. Then, the bias-correction factor for the t[th] estimator is $\hat{z}_t = \Phi^{-1}[H_{B,t}(\hat{h}_{t+1})]$, and the critical values for the individual tests are the $\Phi(2\hat{z}_t + z^{\beta/2})$ and $\Phi(2\hat{z}_t + z^{1-\beta/2})$ quantiles of $H_{B,t}$, where $z^{\beta/2} = \Phi^{-1}(\beta/2)$. Note that if $\hat{h}_{t+1}$ is the median value of the bootstrap distribution, then $\hat{z}_t = 0$ and the uncorrected quantiles of the bootstrap distributions are used for the test regions.

Testing the CDR-PVM with Infinite Holding Horizon

The test of the infinite-holding-horizon CDR-PVM is based on expression (3.10) and requires a time-series model of real dividends (D_t). By setting $\delta_{t+n} = \delta \ \forall$ t, multiplying (3.8) and (3.9) by $(1 - \delta)$, and rearranging we obtain (4.5):

$$(4.5) \quad -T_\Sigma \leq g_t \leq T_\Pi,$$

where

$$(4.6) \quad g_t \equiv \frac{\delta}{(1-\delta)\,P_t}\,[E_t(D_{t+1}) + \sum_{s=2}^{\infty} \delta^{s-1} E_t(\Delta D_{t+s})] - 1$$

and $\Delta D_t \equiv D_t - D_{t-1}$. Expression (4.5) is the testable implication of the infinite-holding-horizon CDR-PVM in the presence of transaction costs. The reason for using (4.5) instead of the CDR version of (3.10) is that the former is more convenient when the real dividend series (D_t) has a stochastic unit root. As reported later, in our specific applications real dividends turn out to have a stochastic unit root.

The test for the infinite-holding-horizon CDR-PVM (4.5) is analogous to that of the one-period-holding-horizon CDR-PVM (3.7). That is, the null hypothesis is analogous to (4.2) such that[4]

[4]The number of lags (p) in (4.7) is the order of the AR(p) model for real dividends.

(4.7) $[\underline{G}_t^{\beta(\alpha)}, \overline{G}_t^{\beta(\alpha)}]$ is a $[1 - \beta(\alpha)]$% confidence interval for g_t, $p \leq t \leq N$.

Then, the infinite-holding-horizon CDR-PVM in the presence of transaction costs can be rejected at the α% significance level if at least one of the $(N - p + 1)$ confidence intervals (4.7) lies entirely below $-T_\Sigma$ or above T_Π. This will happen if and only if conditions (4.8) or (4.9) hold:

$$(4.8) \quad -T_\Sigma > \min[\overline{G}_p^{\beta(\alpha)}, \ldots, \overline{G}_N^{\beta(\alpha)}],$$

$$(4.9) \quad T_\Pi < \max[\underline{G}_p^{\beta(\alpha)}, \ldots, \underline{G}_N^{\beta(\alpha)}].$$

This test is conducted with the aforementioned bootstrap method.

For the infinite-holding horizon case, bootstrapping is even more helpful because the g_t variable is highly nonlinear in the parameters of the time series representation of real dividends. For example, the first-differenced real dividend series analyzed here turn out to be represented by AR(1) processes. That is,

$$(4.10) \quad \Delta D_t = \theta_1 \Delta D_{t-1} + e_t,$$

where θ_1 is a fixed coefficient and e_t is an i.i.d. error term. In this instance, the g_t variable is given by:

$$(4.11) \quad g_t = \frac{\delta}{(1-\delta) P_t} (D_t + \frac{\theta_1}{1-\delta\,\theta_1} \Delta D_t) - 1.$$

Testing for Structural Change

The series analyzed span almost one century and the possibility of structural change is quite real. As mentioned earlier, however, to date the null hypothesis of parameter stability (i.e., no structural change) has not been formally tested. Here, the Wald test statistic recently developed by Andrews is used to test for structural change without specifying the precise time of the likely structural change.[5] Testing for structural change without postulating a specific change point seems appropriate here, because a priori there are numerous change-point candidates (e.g., first world war, second world war, great depression, mid-1970s commodity boom). It is also worth noting that, although Andrews' Wald statistic is designed to test for one-time changes, the test has power against more general forms of parameter instability.

Data

The basic series used in the econometric tests are $P_t^i \equiv p_t^i / CPI_t$ and $D_t^i \equiv d_t^i / CPI_t$, for i = USDA and ISU. Variable p^{USDA} is the per-acre value of cash-rented Iowa farm real estate and d^{USDA} denotes the corresponding gross cash rents minus property taxes calculated by the USDA from survey data.[6] The publication *Major Statistical Series of the U.S. Department of Agriculture: Land Values and Land Use* provides details about the construction of the value and rent series. The p^{ISU} and d^{ISU} series are identical to p^{USDA} and d^{USDA}, respectively, except that the former contain the land values and rents reported by Iowa State University Extension Service whenever these are available. The p^{ISU} and d^{ISU} series are included in the analysis for completeness, because Falk has argued that they are of higher quality than the USDA series. Nominal series are expressed in real terms by deflating the corresponding nominal series by the Consumer Price Index (CPI) (source: *The Statistical History of the United States* and Council of Economic Advisors).

The CDRs for the discounted net rate-of-return series h^{USDA} and h^{ISU} are $\delta^{h^{USDA}}$ = 0.941 and $\delta^{h^{ISU}}$ = 0.941, respectively. Following Falk, CDRs are constructed so that the sample means of the h series are set equal to zero; that is, $\delta^{h^i} \equiv 1/\text{mean}[(P_{t+1}^i + D_{t+1}^i)/P_t^i]$ (i = USDA, ISU). Hence, the two h series have zero means by construction. Standard

[5]Andrews' test for structural change is performed after the unit root test because structural changes tend to bias the latter against rejection of the null hypothesis of unit roots (Enders 1995, p. 243).

[6]The kind assistance of NNN at the USDA in providing the data set and related information is gratefully acknowledged.

deviations of h^{USDA} and h^{ISU} are 0.088 and 0.090, respectively. As well, the minimum (maximum) values of h^{USDA} and h^{ISU} are respectively –0.317 (0.243) and –0.305 (0.225).

CDRs for the g series are constructed in a manner analogous to CDRs for the h series, which yields $\delta^{g^i} \equiv 1/[1 + \text{mean}(D^i_{t+1}/P^i_t)]$ (i = USDA, ISU). The CDRs thus obtained for the g^{USDA} and g^{ISU} series are $\delta^{g^{USDA}} = 0.950$ and $\delta^{g^{ISU}} = 0.950$, respectively.

Price and dividend (i.e., net cash rent) series span the period 1900 through 1994. To our knowledge, the Iowa data are the longest high-quality annual series of farm real estate prices with respective cash rents and property taxes for the U.S.[7] Iowa is the only state for which there are farmland price and rent data dating back to 1900.[8] The Iowa data set is unique because it covers adequately the two major land price cycles that occurred in this century (Figure 1).

Results and Discussion

One-Period Holding Horizon

Results for the trend tests corresponding to the h series are summarized in Table 1. The top half of Table 1 reports results of the test for the null hypothesis of a stochastic trend using the Dickey-Fuller τ_τ statistic. For both h series, the null hypothesis is rejected at the 1% level of significance. Results of the test corresponding to the null hypothesis of no deterministic trend are shown in the bottom half of Table 1. In this instance, the Student-t statistic does not allow us to reject the null hypothesis at any reasonable significance level for any of the h series. Hence, it can be concluded that neither h^{USDA} nor h^{ISU} contain deterministic or stochastic trends.

Estimates of the best-fitting AR(p) models are displayed in Table 2, along with various test statistics for model adequacy. For both h series, the best fit is obtained with an AR(1). The explanatory power of the standard AR(1) is moderate ($R^2 = 0.309$ for h^{USDA} and $R^2 = 0.345$ for h^{ISU}), but the t-statistics on the lagged dependent variable are highly significant in both series (t = 6.41 for h^{USDA} and t = 6.96 for h^{ISU}). Therefore, the

[7]Other studies using long data series are Featherstone and Baker, Hanson and Myers, and Falk. However, their series are respectively 19, 14, and 29 years shorter than the present series.

[8]For other states, farmland price and rent data go back to 1921 at the most.

Table 1. Tests for stochastic and deterministic trends of net discounted rate of return (h) series.

Models: (1) $\Delta h_t = \alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \textit{time} + \alpha_{2+1} \Delta h_{t-1} + \ldots + \alpha_{2+p} \Delta h_{t-p} + e_t$

(2) $h_t = \beta_0 + \beta_2 \textit{time} + \beta_{2+1} h_{t-1} + \ldots + \beta_{2+p} h_{t-p} + e_t$

Null	Evaluation Statistics		Critical value	
Hypothesis	h_t^{USDA}	h_t^{ISU}	5%	1%
$\alpha_1 = 0$	$\tau_\tau = -5.23$	$\tau_\tau = -4.98$	−3.46	−4.06
$\beta_2 = 0$	$t = 0.09$	$t = 0.12$	1.99	2.63

Note: τ_τ and t are the Dickey-Fuller's τ_τ and Student-t test statistics, respectively. The lag length (p) is chosen on the basis of the Schwarz-Bayes Information Criterion. For both USDA and ISU series, p = 0 for model (1) and p = 1 for model (2).

Table 2. Ordinary least squares regressions corresponding to autoregressive models of net discounted rate of return (h) series.

MODEL	TESTS				
	Str. Change	e_t autocorrelation		e_t ARCH	
	(AW)[a]	(Q'(1))[b]	(Q'(2))	(LA(1))[c]	(LA(2))
$h_t^{USDA} = 0.548^{**}\ h_{t-1}^{USDA} + e_t,$ (0.085)[d]	4.22 [> 0.10][e]	0.04 [0.84]	0.46 [0.79]	0.05 [0.82]	0.31 [0.86]
$R^2 = 0.309,\ \hat{\sigma}_e = 0.073$, 93 observations					
$h_t^{ISU} = 0.580^{**}\ h_{t-1}^{ISU} + e_t,$ (0.084)	3.36 [> 0.10]	0.20 [0.66]	0.27 [0.88]	1.77 [0.18]	5.32 [0.07]
$R^2 = 0.345,\ \hat{\sigma}_e = 0.073$, 93 observations					

* (**) Significantly different from zero at the 5 (1) percent level of significance based on the two-tailed t-statistic.
[a]AW denotes Andrews' Wald-like test for structural change.
[b]Q'(i) is the Ljung-Box portmanteau test or modified-Q statistic for i-order autocorrelation (Ljung and Box).
[c]LMA(i) is the Lagrange multiplier test for i-order autoregressive conditional heteroscedasticity (Engle).
[d]Numbers between parentheses below coefficient estimates denote the respective standard deviations.
[e]Numbers between brackets below test statistics denote the corresponding *p*-values.

frictionless CDR-PVM is strongly rejected for farmland prices. This finding is consistent with the results reported in the existing literature.

Results of model specification tests are reported in the columns under the "tests" heading. The modified-Q statistic does not provide indication of first- or second-order autocorrelation in the errors. As well, Engle's Lagrange multiplier test for first- and second-order ARCH does not evince ARCH problems. More importantly, the Andrews test for structural change does not allow us to reject the null hypothesis of no structural change in any of the h series analyzed for the period under study. Finally, the residuals for each of the models were used to conduct Jarque-Bera (i.e., skewness-kurtosis) and goodness-of-fit (Mittelhammer, p.656) tests for normality of the error distribution. Three outliers for the USDA series lead to a Jarque-Bera test statistic of 16.94 (χ_2^2 *p*-value is 0.002) whereas the ISU series has a test statistic of 3.37, which does not lead to rejection of normality at the usual significance levels. Alternatively, the goodness-of-fit tests reject the normality hypothesis for the USDA and ISU series with respective *p*-values of 0.0058 and 0.0004. The mixed sample evidence for the normality hypothesis provides further support for our use of the bootstrap testing procedure.

Figure 2 summarizes the results of the tests of the CDR-PVM in the presence of transaction costs. The curves drawn in Figure 2 represent the minimum level of total transaction costs necessary to avoid rejecting the CDR-PVM at the test size specified in the horizontal axis.[9] For example, the vertical line depicted at the 5% overall test size intersects the USDA (ISU) curve at a total transaction cost level of 6.15% (7.34%). Therefore, the null hypothesis that the CDR-PVM holds for transactions costs of 6.15% (7.34%) or more cannot be rejected at the 5% significance level for the USDA (ISU) farmland series. Given that transaction costs in farmland market are estimated to be 7.50% at a minimum (Wunderlich), we can conclude that the data do not provide enough evidence to reject the CDR-PVM for farmland. The horizontal line drawn at total transaction costs of 7.50% indicates that the null hypothesis that the CDR-PVM holds for

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[9]Due to lack of better knowledge, in Figure 2 buying and selling transaction costs are assumed to be the same (i.e., $T_\Pi = T_\Sigma = T$, and total transaction costs equal 2 T). Graphs analogous to Figure 2 can be readily constructed for different levels of buying and selling transaction costs.

Figure 2. Smallest total transaction costs consistent with the CDR-PVM null hypothesis (one-period-holding-horizon model).

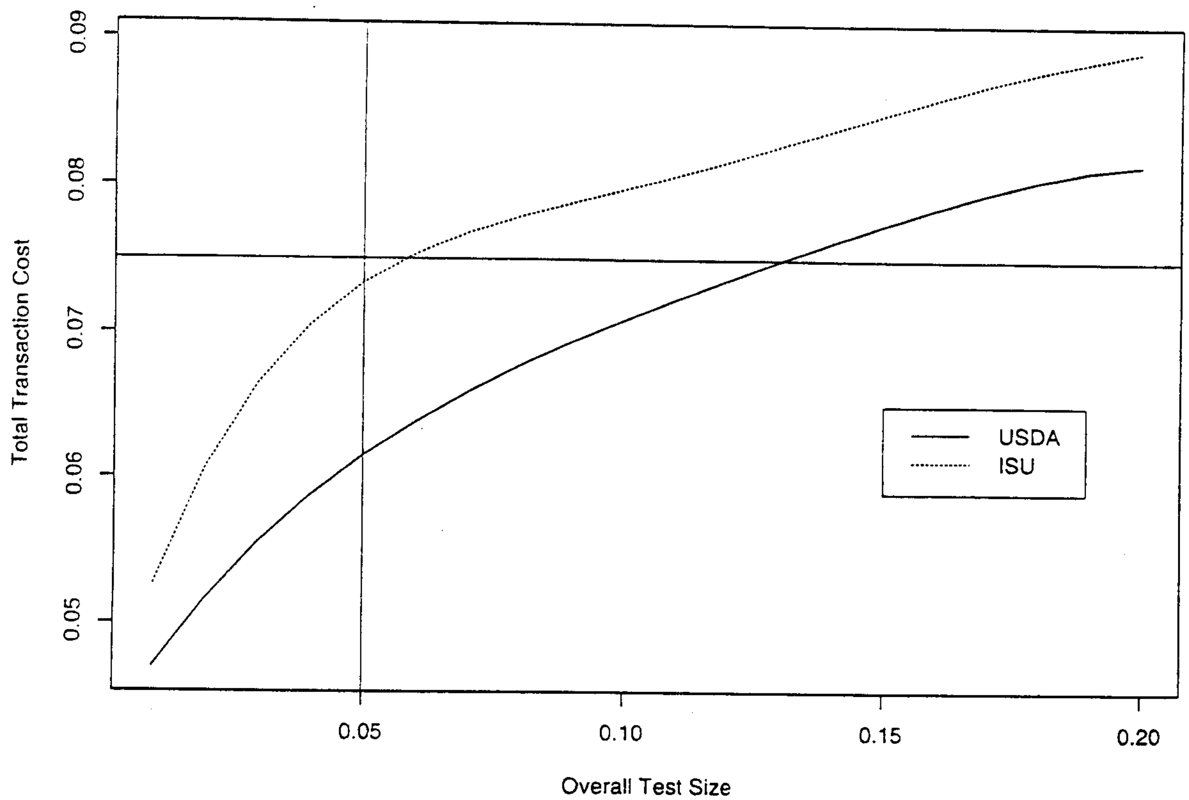
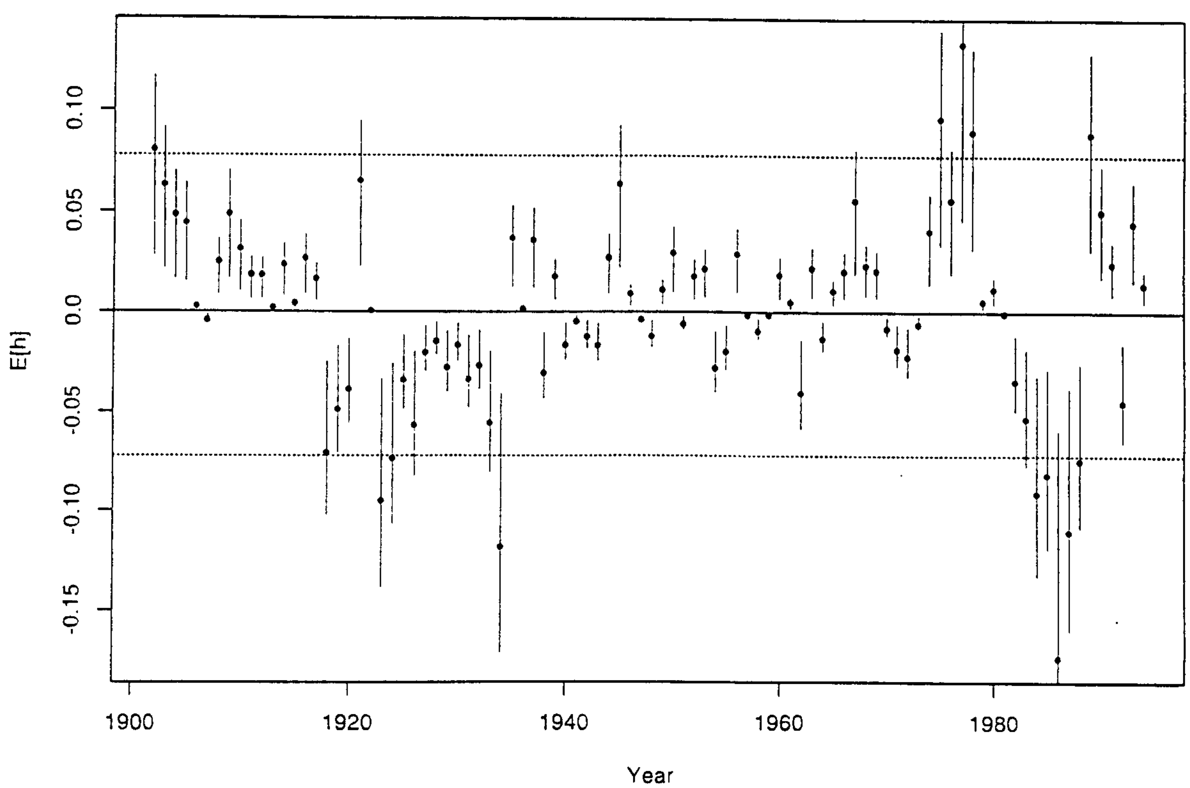


Figure 3. Point estimates and 95% conf. intervals for conditional expectations of h^{USDA}.



total transaction costs of 7.50% can be rejected at significance levels of 13.04% (5.80%) or greater for the USDA (ISU) series.

Point estimates along with bias-corrected confidence intervals for the conditional expectations of h^{USDA} are plotted in Figure 3. It can be observed that most of the confidence intervals lie entirely above or below the horizontal line drawn at $\lambda^L = \lambda^U = 0$. However, none of the confidence intervals lies entirely outside of the horizontal band ranging from $\lambda^L = -7.23\%$ through $\lambda^U = 7.79\%$ (i.e., the bounds corresponding to total transaction costs of 7.50%, assuming $T_\Pi = T_\Sigma = 3.75\%$). These results are consistent with our earlier rejection of the frictionless CDR-PVM and nonrejection of the CDR-PVM in the presence of transaction costs of 7.50%. It is also worth noting that the confidence intervals are asymmetric around the conditional mean estimates due to the bias-correction procedure. Results for the h^{ISU} series are similar and are omitted to save space.

Infinite Holding Horizon

Testing the infinite-holding-horizon CDR-PVM requires a time series model of real dividends (D). The first step in doing so is to test for unit roots.

Table 3 reports the results from the unit-root tests for the real dividend series. The first row reveals that the null hypothesis of two unit-roots ($\gamma_1 = 0$) is strongly rejected for both series. In contrast, the results in the second row show that the null hypothesis of one unit root in the augmented model (2) ($\alpha_1 = 0$) cannot be rejected at standard levels of significance for any of the series. Following the procedure outlined in Enders (1995), the next step is to test whether too many regressors are included in the augmented model (2). According to the results shown in the third row, the trend term is not significantly different from zero in any of the series. Hence, regressions are re-estimated without the time trend (i.e., model (3)). Using model (3), the fourth row reveals that the null hypothesis of a unit-root ($\eta_1 = 0$) cannot be rejected at usual significance levels for any of the series. Because of the failure to reject the unit-root hypothesis, the next step consists of testing for the significance of the constant term in model (3). The test statistics in the fifth row do not allow us to reject the null hypothesis of no constant term for any of the series. Therefore, regressions are re-estimated without the constant term (i.e., model (4)). The

Table 3. Tests for stochastic and deterministic trends of real dividend (D) series.

Models: (1) $\Delta(\Delta D_t) = \gamma_0 + \gamma_1 \Delta D_{t-1} + \gamma_2\, time + \gamma_{2+1} \Delta(\Delta D_{t-1}) + \ldots + \gamma_{2+p} \Delta(\Delta D_{t-p}) + e_t$

(2) $\Delta D_t = \alpha_0 + \alpha_1 D_{t-1} + \alpha_2\, time + \alpha_{2+1} \Delta D_{t-1} + \ldots + \alpha_{2+p} \Delta D_{t-p} + e_t$

(3) $\Delta D_t = \eta_0 + \eta_1 D_{t-1} + \eta_{2+1} \Delta D_{t-1} + \ldots + \eta_{2+p} \Delta D_{t-p} + e_t$

(4) $\Delta D_t = \xi_1 D_{t-1} + \xi_{2+1} \Delta D_{t-1} + \ldots + \xi_{2+p} \Delta D_{t-p} + e_t$

Null	Evaluation Statistics		Critical value	
Hypothesis	D_t^{USDA}	D_t^{ISU}	5%	1%
$\gamma_1 = 0$	$\tau_\tau = -5.84$	$\tau_\tau = -6.34$	−3.46	−4.06
$\alpha_1 = 0$	$\tau_\tau = -1.92$	$\tau_\tau = -1.86$	−3.46	−4.06
$\alpha_2 = 0$	$\tau_{\beta\tau} = 1.10$	$\tau_{\beta\tau} = 1.08$	2.79	3.53
$\eta_1 = 0$	$\tau_\mu = -1.58$	$\tau_\mu = -1.52$	−2.89	−3.52
$\eta_0 = 0$	$\tau_{\alpha\tau} = 1.54$	$\tau_{\alpha\tau} = 1.49$	2.54	3.22
$\xi_1 = 0$	$\tau = -0.41$	$\tau = -0.36$	−1.95	−2.60

Note: τ_τ, $\tau_{\beta\tau}$, τ_μ, $\tau_{\alpha\tau}$, and τ are the Dickey-Fuller's τ_τ, $\tau_{\beta\tau}$, τ_μ, $\tau_{\alpha\tau}$, and τ test statistics, respectively. The lag length (p) is chosen on the basis of the Schwarz-Bayes Information Criterion. For both USDA and ISU series, p = 0 for model (1) and p = 1 for models (2) through (4).

last row reveals that the hypothesis of a unit-root still cannot be rejected for any of the series at any typical significance level. Therefore, we can conclude that both real dividend series contain one stochastic unit root.

The selected AR(p) models for the first-differenced real dividend series are displayed in Table 4. Both farmland dividends are adequately represented by an AR(1). The explanatory power is relativey low ($R^2 < 0.20$). Andrews' test provides no evidence of structural change for any of the real dividend series analyzed. Based on the autocorrelation and ARCH tests, both AR(1) models seem well specified. However, Jarque-Bera tests strongly reject the null hypothesis of normality of residuals in both instances. This finding lends support for using bootstrap methods to test the CDR-PVM.

For the infinite holding horizon, the smallest transaction costs consistent with the CDR-PVM hypothesis are depicted in Figure 4. At the 5% level of significance, total transaction costs would need to be at least 96.05% and 97.08% for the ISU and the USDA series, respectively. Such transaction costs are far above the normal cost levels observed in farmland transactions. It can therefore be concluded that the data strongly rejects the infinite holding-horizon CDR-PVM.

Figure 5 displays the point estimates and the 95% confidence intervals corresponding to g^{USDA}. Most of the 95% confidence intervals do not overlap with the horizontal line drawn at zero, which indicates that the frictionless CDR-PVM does not hold for the USDA series. Furthermore, Figure 5 indicates that the CDR-PVM is also strongly rejected even after considering typical levels of transaction costs in farmland markets. For example, most of the 95% confidence intervals in Figure 5 lie entirely outside of the band $T_\Sigma = -0.05$ and $T_\Pi = 0.05$, which indicates that the CDR-PVM of farmland is strongly rejected for total transaction costs as high as 10%. Results for g^{ISU} are quite similar and are omitted in the interest of space.

The infinite-holding horizon assumption seems the most reasonable for farmland, as it is well-known that farmland changes hand very infrequently. According to Rogers and Wunderlich, the average turnover of farmland is approximately 28 years. Therefore, our test results strongly reject the CDR-PVM for farmland, even after accounting for the typical transaction costs that characterize farmland markets.

Table 4. Ordinary least squares regressions corresponding to autoregressive models of first differenced real dividend (ΔD) series.

MODEL	TESTS				
	Str. Change	e_t autocorrelation		e_t ARCH	
	(AW)[a]	(Q'(1))[b]	(Q'(2))	(LA(1))[c]	(LA(2))
$\Delta D_t^{USDA} = 0.448^{**}\ \Delta D_{t-1}^{USDA} + e_t,$	5.84	0.03	0.23	2.23	3.57
(0.094)[d]	[> 0.10][e]	[0.86]	[0.89]	[0.14]	[0.17]
$R^2 = 0.199,\ \hat{\sigma}_e = 0.013$, 93 observations					
$\Delta D_t^{ISU} = 0.381^{**}\ \Delta D_{t-1}^{ISU} + e_t,$	5.82	0.01	0.03	3.78	4.28
(0.097)	[> 0.10]	[0.92]	[0.99]	[0.05]	[0.12]
$R^2 = 0.143,\ \hat{\sigma}_e = 0.014$, 93 observations					

* (**) Significantly different from zero at the 5 (1) percent level of significance based on the two-tailed t-statistic.
[a]AW denotes Andrews' Wald-like test for structural change.
[b]Q'(i) is the Ljung-Box portmanteau test or modified-Q statistic for i-order autocorrelation (Ljung and Box).
[c]LMA(i) is the Lagrange multiplier test for i-order autoregressive conditional heteroscedasticity (Engle).
[d]Numbers between parentheses below coefficient estimates denote the respective standard deviations.
[e]Numbers between brackets below test statistics denote the corresponding *p*-values.

Figure 4. Smallest total transaction costs consistent with the CDR-PVM null hypothesis (infinite-holding-horizon model).

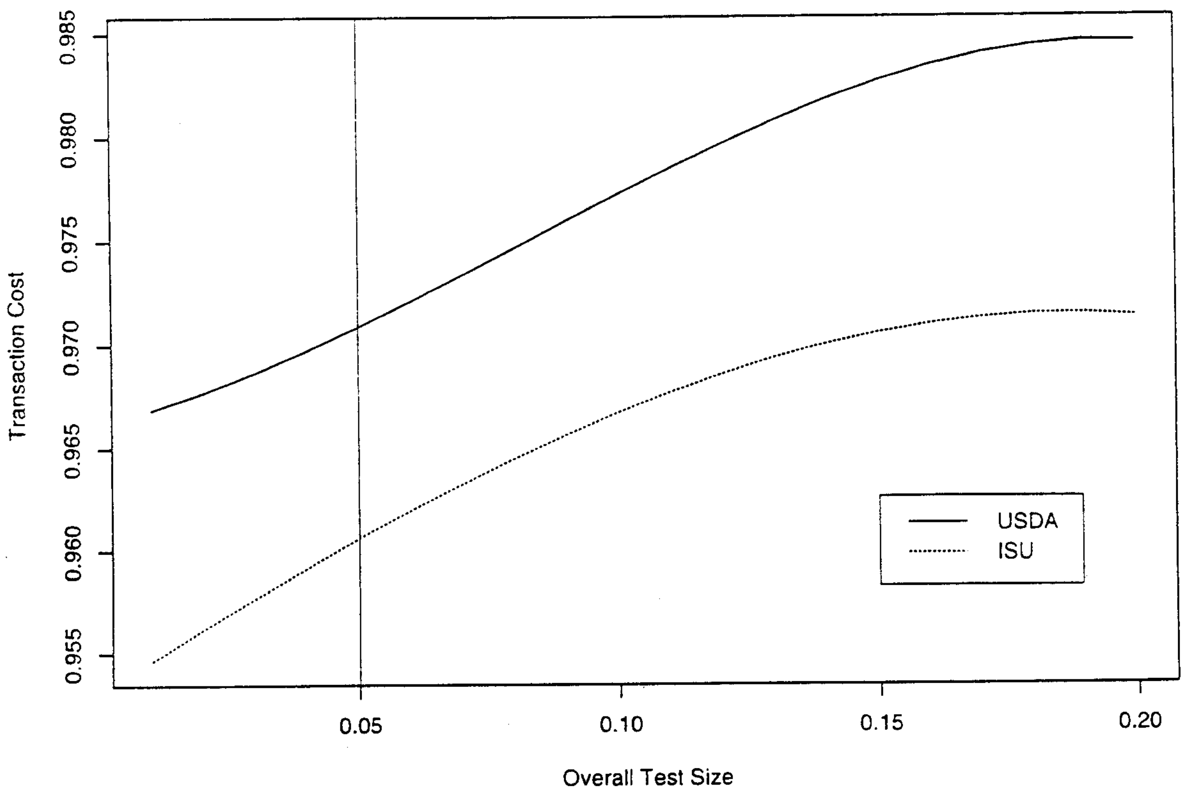
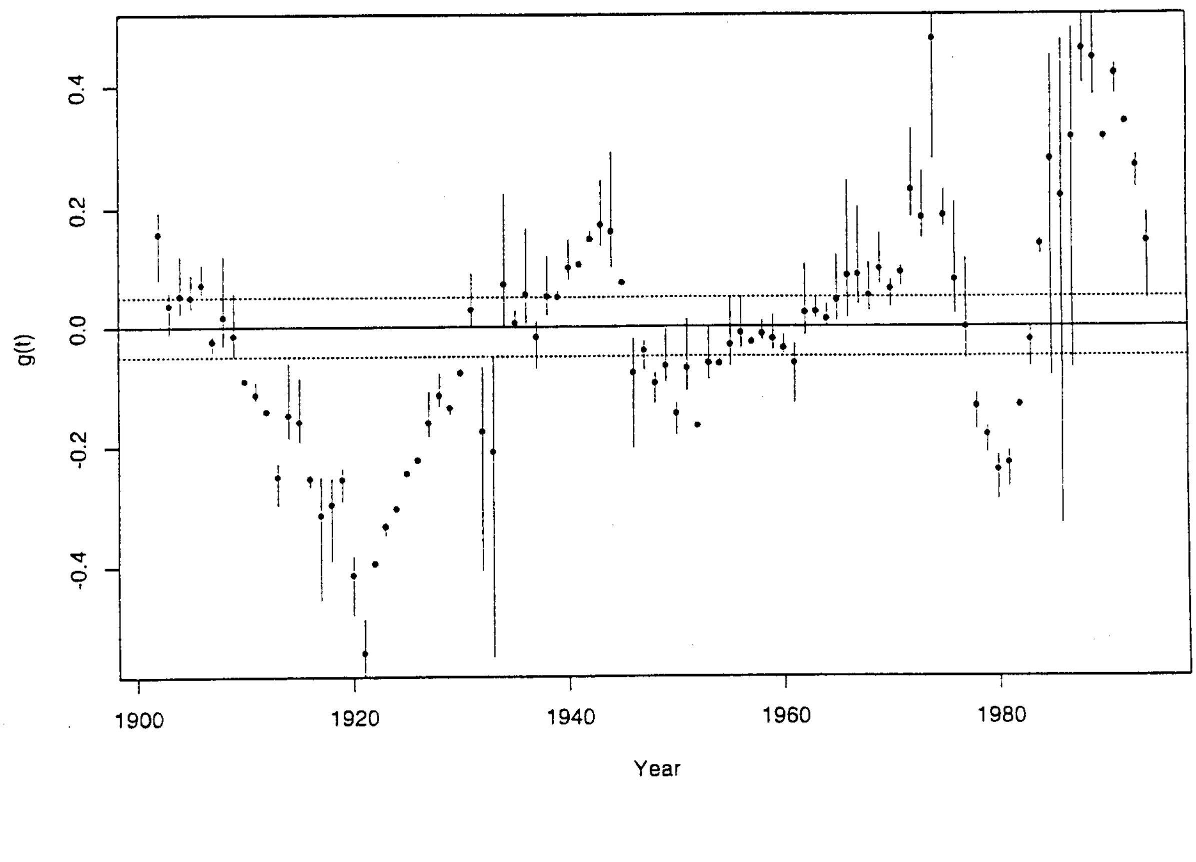


Figure 5. Point estimates and 95% confidence intervals for g^{USDA}.



Concluding Remarks

Despite the voluminous work on land prices, the existing literature has devoted little attention to the large transaction costs involved in transfers of farmland ownership for farmland price behavior. The present study uses the longest known time series available on U.S. farmland price and rents to analyze empirically the implications of transaction costs for farmland pricing. These data are valuable because they track the two largest boom-bust cycles in farmland prices that occurred in this century.

As well, no study has previously tested for the presence of structural change in the returns to farmland. This is somewhat paradoxical, given that the series used to analyze farmland price behavior usually extend over many years. Our long series are ideally suited to study the possible presence of structural changes. Interestingly, our data do not provide significant evidence of structural change.

Two alternative bootstrapping tests of the CDR-PVM in the presence of transaction costs are developed. Such tests are associated with the polar cases of a one-period holding horizon and an infinite-period holding horizon. According to the results of the one-period holding horizon test, the CDR-PVM of farmland is strongly rejected assuming no transaction costs, but cannot be rejected in the presence of the typical transaction costs involved in the transfer of farmland ownership.

In contrast, the infinite-holding horizon test strongly rejects the CDR-PVM for farmland at any reasonable level of transaction costs. Given the conservatism of the one-period holding horizon test and the fact that land is typically held for many years, the infinite-holding-horizon test results seem the most relevant for farmland.

In summary, the present study indicates that farmland price behavior is inconsistent with the CDR-PVM, even after considering the typical transaction costs involved in farmland markets. The analysis also reveals that the failure of the CDR-PVM of farmland cannot be attributed to structural change. Therefore, the failure of the CDR-PVM must be explained by reasons other than transaction costs or structural change. Exploring such reasons seem important and promising avenues for further research.

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