Rents and Land Prices in Japan: A Panel Cointegration Approach

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Rents and Land Prices in Japan: A Panel Cointegration Approach

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ABSTRACT. *The Japanese farmland market is strongly regulated, although partial deregulation and decentralization are evident. This paper examines the relationship between farmland rents and prices in Japan using recent panel cointegration methods, which admit structural breaks. Results show the presence of a cointegrating relationship with significant breaks that increased the rent/price ratio by 9% in 1967 and by 15% in 1980; prices cause rents, which supports an institutional rent-formation hypothesis; and the farmland market is inefficient.* (JEL C51, Q15)

I. INTRODUCTION

The literature on the relationship between farmland prices and rents is extensive; examples include work by Featherstone and Baker (1987), Falk (1991), Lloyd, Rayner, and Orme (1991), Lloyd (1994), and Lence and Miller (1999). In many studies, the present-value model (PVM) provides the hypothesis that prices are determined by rents. In Japan, however, the farmland market is strongly regulated with rents being influenced through a process of institutional governance guided by land prices. Shigeto, Hubbard, and Dawson (2008) examined the farmland rent-price nexus in Japan using national data and the cointegration procedure of Johansen, Mosconi, and Nielsen (2000) where structural breaks model rent revisions in 1967 and 1980. They concluded that prices cause rents, which supports an institutional rent-formation hypothesis rather than the PVM, and that the rent/price ratio increased significantly in 1980, by 21%, but not in 1967. Perhaps surprisingly, they find that the price-rent elasticity is unity, which supports the notion that the Japanese farmland market is efficient. This conclusion runs counter to the claims of some commentators who believe that this market is inefficient and distorted, owing to government interference and the distinctive nature of farmers' behavior (e.g., Honma 1994; Egaitsu and Shogenji 1995; Kusakari 1998; Godo 1998, 2006, 2007).

Most multivariate analyses of the farmland rent-price nexus use a time series of moderate length and conventional cointegration tests. Such tests are subject to two criticisms: first, they tend to underreject in the presence of structural breaks; and second, and particularly pertinent to Shigeto, Hubbard, and Dawson (2008), they have poor size and power properties. Gutierrez, Westerlund, and Erickson (2007) address these criticisms by using panel data for 31 U.S. states for 1960–2000 and the method of Westerlund (2006) to seek a panel cointegrating relationship between rents and prices with unknown breaks. Their results show no evidence of cointegration in the absence of structural

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breaks, but when breaks are admitted, rents and prices are cointegrated.

The conclusions of Shigeto, Hubbard, and Dawson (2008) therefore should be treated with caution. This paper reexamines the rent-price relationship in the Japanese farmland market using a richer, more disaggregated panel dataset for nine regions for 1955–2000. In seeking more robust results than those of Shigeto, Hubbard, and Dawson, especially in relation to their result that the Japanese farmland market is efficient, we apply the panel cointegration method of Westerlund (2006), which admits structural breaks to estimate both regional and aggregate relationships, and we examine the effects of the rent revisions in 1967 and 1980. We also use the panel causality tests of Canning and Pedroni (2008) to test between the PVM and institutional models; this appears to be the first use of this test in the agricultural economics literature.

II. THE FARMLAND MARKET IN JAPAN

The farmland market in Japan is subject to a system of strict controls, although "deregulation and decentralization" are evident (Godo 2007). Land reform began immediately after the Second World War as part of the transformation from feudalism to democracy. The main pillars of land reform were the reduction of rent and conversion of rent-in-kind to money rent, the creation of "owner-cultivators," and the democratization of agricultural land committees (Koppel and Kim 1993). The Agricultural Land Law (1952) restricted farmland ownership to individual owner-cultivators rather than to business corporations because of concerns over speculative possession,[1] and the maximum level of farm rent was controlled. Notably, the maximum rent level was revised in 1967, and direct rent control was abolished in 1970; but, as a transitional measure, controls were applied to existing tenanted land for a further 10 years. This caused a marked increase in rents at the beginning of the 1980s. Following these changes, the control system evolved into a standard system that gave local Chambers of Agriculture a role in setting rents and advising on reductions where they were considered too high.

By contrast, farmland prices in Japan have not been directly controlled. Instead, zoning has been applied under various laws. In exclusively agricultural areas, land is not allowed to convert to nonagricultural use, so as to prevent speculative price movements. Godo (1998, 2007) criticizes the effectiveness of the regulations and argues that the farmland market is distorted because farmers rarely sell or lease land and may even opt for land abandonment in expectation of windfall gains from buy-outs by public authorities, which often require land for nonagricultural use.[2] These characteristics lead Shigeto, Hubbard, and Dawson (2008) to conclude that an institutional-based rent-formation model is more appropriate than the PVM. We investigate this idea further using regional data.

III. EMPIRICAL MODEL AND METHOD

Following inter alia Burt (1986), Featherstone and Baker (1987), and Falk (1991), the PVM of the farmland market can be represented as

$$P_t = \alpha \sum_{j=1}^{\infty} \alpha^j E_t[R_{t+j}], \qquad [1]$$

where P_t and R_t are the equilibrium price and rent in period t; α is a constant discount factor equal to $1/(1 + i)$, where i is the real discount rate; and E_t is the conditional expectation based on information available at time t. Here, price equals the expected present value of future rents, and there is a long-run relationship between the real price of farmland and its real return. Denoting p_t and r_t as the long-run equilibrium price and rent in logarithms, [1] becomes

[1] Agricultural production corporations have been able to own farmland since 1962.

[2] Regulations prohibit land abandonment, but penalties are ineffective.

$$p_t = \beta_0 + \beta_1 r_t, \quad [2]$$

where $\beta_1 = 1$, that is, the market is efficient (Lloyd 1994), and $\beta_0 = \log(1/i)$. Shigeto, Hubbard, and Dawson (2008) propose an alternative "institutional model" where rent is set in accordance with price by a process of institutional governance, that is,

$$r_t = \beta_0' + \beta_1' p_t. \quad [3]$$

Here, rent is a mark-down on price. Policy changes are modeled by differential intercepts in [2] and [3].

There are three objectives of our empirical investigation. First, we examine the possible existence of cointegration, thereby inquiring whether a long-run relationship exists between rents and prices. Conditionally, we then estimate this relationship and examine both the impact of the rent revisions in 1967 and 1980, and market efficiency. Finally, we test for causality to identify whether the PVM in [2] or the institutional model in [3] is more appropriate.

Westerlund (2006) develops a Lagrange multiplier (LM) statistic to test the null of panel cointegration where structural breaks are permitted in each individual relationship in both null and alternative hypotheses. Adapting this test to the case here and using the institutional model in [3] to illustrate, the data-generating process for r_{it} is

$$r_{it} = \mathbf{z}_{it}'\boldsymbol{\delta}_{iq} + p_{it}\beta_i + \varepsilon_{it}, \quad [4]$$

$$\varepsilon_{it} = \gamma_{it} + \upsilon_{it}, \quad [5]$$

$$\gamma_{it} = \gamma_{it-1} + \phi_i \upsilon_{it}, \quad [6]$$

where rents, r_{it}, and prices, p_{it}, are $I(1)$ processes; the individual regions are $i = 1,\ldots,N$ and time periods $t = 1,\ldots,T$; $\mathbf{z}_{it}$ is a vector of deterministic variables that may include constants, trends, and segmented constants and trends specific to each region i; $\boldsymbol{\delta}_{iq}$ for $q = 1, 2, 3$ is a vector of estimated parameters where there are two breaks ($m = 2$) (and three regimes) in cross-section i, which are located at dates $T_{i1} = 1967$ and $T_{i2} = 1980$; β_i is the cointegrating parameter; the errors, ε_{it}, are generated as the sum of a random walk, γ_{it}, and a stationary stochastic process, $\upsilon_{it} \sim iid(0,\sigma^2_{\upsilon_i})$; and γ_{i0} can be assumed to be zero if constants are included in $\mathbf{z}_{it}$. The null is that all regions in the panel are cointegrated, that is, $\phi_i = 0$ for all $i = 1,\ldots,N$ against the alternative that $\phi_i \neq 0$ for some i.[3] When $\phi_i = 0$, γ_{it} in [5] and [6] vanishes (assuming $\gamma_{i0} = 0$), $\varepsilon_{it} = \upsilon_{it}$, and r_{it} and p_{it} are cointegrated since υ_{it} is stationary. The panel LM-statistic is the cross-section average of the three regime-specific KPSS-statistics (Kwiatkowski et al. 1992):

$$\mathrm{LM} = \frac{1}{N}\sum_{i=1}^{N}\sum_{q=1}^{m+1}\sum_{t=T_{iq-1}+1}^{T_{iq}} \frac{1}{(T_{iq}-T_{iq-1})^2}\frac{S_{it}^2}{\hat{\sigma}_i^2}, \quad [7]$$

where $S_{it} = \sum_{s=T_{iq-1}+1}^{t} \hat{\varepsilon}_{is}$ is the partial sum of efficient estimates of the residuals in [4], and $\hat{\sigma}_i^2$ is a consistent estimate of the long-run variance of ε_{it}. The corresponding standardized statistic is distributed as standard normal under the assumption of cross-independency:

$$\text{Z-LM} = \frac{\sqrt{N}(\mathrm{LM} - \bar{\Theta})}{\sqrt{\bar{\Sigma}}} \sim N(0,1), \quad [8]$$

where $\bar{\Theta}$ and $\bar{\Sigma}$ are the average of the mean and variance of the limiting distribution of the LM-statistic in [7]. Response surface moments are obtained by Westerlund using Monte Carlo simulations of the limiting distribution, which depend on the deterministic specification of the model, and the number of regressors (but not on the location of the breakpoints). The Z-LM-statistic is compared with the right tail of the normal distribution.

[3] Note that ϕ_i can be zero for some regions, and it is not required that $\phi_i = \phi \neq 0$.

Pedroni (2001) and Westerlund (2006) consider two panel estimators for obtaining the parameters in [4]: dynamic ordinary least squares (DOLS) is a parametric method where lags are explicitly estimated, and fully modified ordinary least squares (FMOLS) is a nonparametric method to deal with serial correlation that uses a heteroskedasticity- and autocorrelation-consistent estimator of the long-run covariance matrix. Both correct for OLS bias induced by endogeneity. DOLS or FMOLS can be used to provide within- or between-group estimates. Pedroni (2001, 2002) argues that the between-group (or group mean) estimator is preferred for two reasons: first, it has relatively minor size distortions in small samples; and second, the t-statistics permit more flexible alternative hypotheses and, in particular, $\beta_i \neq \beta_0$ so that β_i need not be the same for all N regions under the alternative. The evidence for preferring DOLS or FMOLS is not so clear and estimates tend to be similar.

We test for causality in the panel following Canning and Pedroni (2008). If cointegration exists between r_{it} and p_{it}, the relationship in [4] can be represented by a dynamic error-correction model (ECM):

$$\Delta r_{it} = \alpha_i^1 \hat{\varepsilon}_{it-1} + \sum_{j=1}^{k} \varphi_{ij}^{11} \Delta r_{it-j} + \sum_{j=1}^{k} \varphi_{ij}^{12} \Delta p_{it-j} + \sum_{q=1}^{m} \tau_i^{1q} I_{it}^{q} + \omega_i^1 + v_{it}^1, \qquad [9]$$

$$\Delta p_{it} = \alpha_i^2 \hat{\varepsilon}_{it-1} + \sum_{j=1}^{k} \varphi_{ij}^{21} \Delta r_{it-j} + \sum_{j=1}^{k} \varphi_{ij}^{22} \Delta p_{it-j} + \sum_{q=1}^{m} \tau_i^{2q} I_{it}^{q} + \omega_i^2 + v_{it}^2. \qquad [10]$$

In [9] for example, $\hat{\varepsilon}_{it-1}$ is the estimated long-run disequilibrium in the previous period, that is, the lagged residuals from [4]; α_i^1 is the error-correction term that gives the reaction of r_{it} to bring the system back to long-run equilibrium; I_{it}^{q} provides the impulse dummies that equal unity when $t = T_{i1} + 1$ ($t = 1968$) and $t = T_{i2} + 1$ ($t = 1981$) to provide consistency with the breaks in the long-run cointegrating relationship in [4];[4] and the ω_i^1 terms are region-specific constants. From the Granger representation theorem (Engle and Granger 1987), causality must exist at least in one direction if a long-run relationship exists between r_{it} and p_{it}, and at least one of the error-correction terms in [9] and [10] must be nonzero.

Consider testing for noncausality for each region. The ECMs in [9] and [10] are estimated separately for each region where the number of region-specific lags, k, with a maximum of $k = 4$ is determined by the Schwartz Bayesian criterion. To illustrate, consider testing for noncausality from p_{it} to r_{it} in [9]. The joint null of no short- or long-run causality for each region is

$$H_0^1: \varphi_{ij}^{12} = 0 \quad \text{and} \quad \alpha_i^1 = 0$$
$$\text{for each } i = 1, \ldots, N \text{ and } j = 1, \ldots, k, \qquad [11]$$

and $F \sim F_{k+1}$. The null of no long-run causality from p_{it} to r_{it} for each region is

$$H_0^2: \alpha_i^1 = 0 \quad \text{for each } i = 1, \ldots, N, \qquad [12]$$

and $t \sim t_{N-2k-m-2}$. While the tests in [11] and [12] are interesting, more general insights are provided by the corresponding panel tests. Again using the same example, we estimate the (single) heterogeneous ECM in [9] for all $i = 1, \ldots, N$ regions where the number of lags, k, are determined from region-specific ECMs as before. The joint null of no short- or long-run causality in the panel is

$$H_0^3: \varphi_{ij}^{12} = 0 \quad \text{and} \quad \alpha_i^1 = 0$$
$$\text{for all } i = 1, \ldots, N \text{ and all } j = 1, \ldots, k, \qquad [13]$$

and the log-likelihood ratio, $\text{LLR} \sim \chi^2_{R+N}$, where R is the total number of lagged Δp_i-terms in all N-equations in [9]. The null of no long-run causality from r_{it} to p_{it} in the

[4] If a segmented trend is specified in the long run, step dummies are included in the short run.

panel is

$$H_0^4: \ \alpha_i^1 = 0 \quad \text{for all } i = 1, \ldots, N, \tag{14}$$

and LLR $\sim \chi_N^2$. The tests in [11] and [13] are Granger-causality tests since they test the null that r_{it} evolves exogenously with respect to p_{it} at all noncontemporaneous time horizons. Evidence in favor of causality from prices to rents supports the institutional model, while causality from rents to prices supports the PVM.

Canning and Pedroni (2008) also test the "pervasiveness" of long-run casual effects in a panel using the group mean test, which is based on the average of regional adjustment coefficients in [9] and [10]. In [9] for example, this average is $\bar{\alpha}^1 = N^{-1} \sum_{i=1}^{N} \alpha_i^1$, the group mean panel t-statistic is $\bar{t}(\alpha^1) = N^{-1} \sum_{i=1}^{N} t(\alpha_i^1)$, and the null of no long-run causality is

$$H_0^5: \ \bar{\alpha}^1 = 0, \tag{15}$$

where $\bar{t}(\alpha^1) \sim N(0,1)$; rejection in favor of the alternative is at either tail. A disadvantage of the group mean test is that while a null average long-run effect implies no causality, it may be a consequence of heterogeneous long-run coefficients with positive and negative values that cancel each other out. To rule this out, Canning and Pedroni propose a homogeneity test. Again using [9], this is a Wald statistic: $W = \sum_{i=1}^{N} \sigma_{\alpha_i}^{-2} (\alpha_i^1 - \bar{\alpha}^1)^2$, where $\bar{\alpha}^1$ represents the estimated group mean estimates of the N error-correction terms and $\sigma_{\alpha_i}^{-2}$ is the inverse of its sample variance. Under the null of parameter homogeneity across regions, $W \sim \chi_N^2$.

IV. DATA AND RESULTS

The dataset comprises annual farmland rents and prices for 1955–2000 for nine regions in Japan, namely, Hokkaido, Tohoku, Kanto, Hokushin, Tokai, Kinki, Chugoku, Shikoku, and Kyushu. They relate to the average price (yen/are[5]) of "good" paddy and "good" vegetable (including grazing) fields.[6] Real prices and rents are calculated using the GDP deflator. The data are shown in Figure 1, and vertical lines indicate the rent revisions in 1967 and 1980. In most regions, there appear to be breaks in rents in 1967 and 1980: average rents across all regions increased by 40% in 1968, and by 13% in 1981 and 34% in 1982. Breaks in prices are not so evident. When expressed as a rent/price ratio, as shown in Figure 2, a noticeable break is evident in 1980, after which the ratios increase substantially.[7]

We test for cointegration between rents and prices by applying the panel test of Westerlund (2006) with breaks in 1967 and 1980. Deterministic trends are not evident, and we model level breaks only. Initially, we use the model in [4], that is,

$$r_{it} = \mu_1 + \mu_2 + \mu_3 + \beta_i p_{it} + \varepsilon_{it}$$
$$\text{for } i = 1, \ldots, N \text{ and } t = 1, \ldots, T, \tag{16}$$

where μ_1, μ_2, and μ_3 are the regime-specific intercepts for 1955–1967, 1968–1980, and 1981–2000.[8]

First, we test for cointegration between rents and prices in each region using a univariate Z-LM-statistic, which is based on [8] before averaging across regions. This is compared with bootstrapped critical values obtained from 5,000 replications following the sieve approach of Westerlund and Edgerton (2007). The results are shown in Table 1.[9] The tests imply nonrejection of

[5] 1 are = 1/100 of 1 hectare = 0.02471 acres = 100 m^2.

[6] The data (Japan Real Estate Institute 2003) are collected from surveys undertaken by the Chambers of Agriculture, to which all local land transactions have to be reported. The chambers exclude outliers that look odd in comparison to the local average, which goes some way toward allaying fears that exceptionally high prices due to speculation may undermine our empirical estimation.

[7] The vertical lines in Figures 1 and 2 are drawn at 1968 and 1981, when the rent revisions became effective.

[8] A model that includes breaks in the trends shows that they are insignificant.

[9] We are grateful to Joakim Westerlund, who provided GAUSS code for the panel cointegration test and bootstrapping.

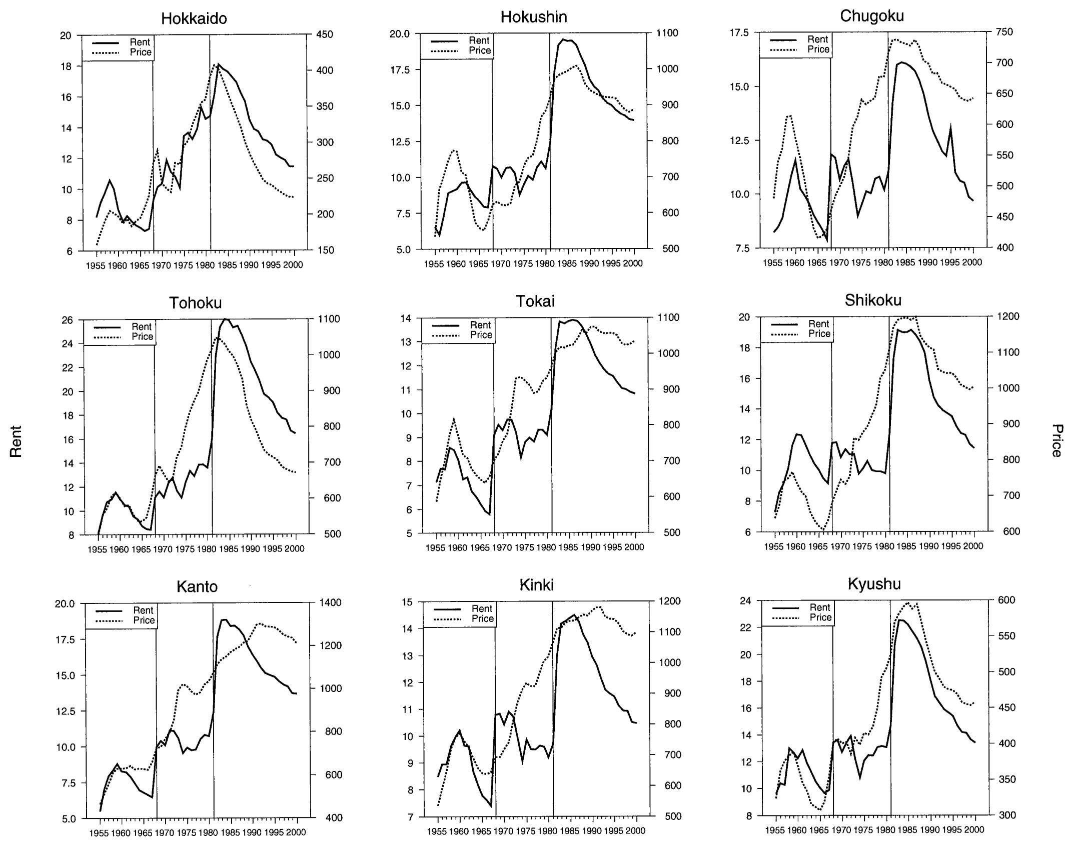

FIGURE 1
FARMLAND RENTS AND PRICES (YEN/10,000 ARE), 1955–2000

the null of cointegration in all regions except for Kantu and Tokai. The panel Z-LM-statistic in [8] does not reject the null of cointegration at the 5% significance level when compared with the bootstrapped distribution, which allows for cross-correlation across regions, and we conclude that cointegration exists between rents and prices.

FMOLS is now used to estimate the cointegrated relationship in [16], and the results are shown in Table 2.[10] First, we test for equality between the constants, and the results are shown in Table 3. For the individual regions, the nulls that the constants between 1955–1967 and 1968–1980 are equal, that is $\mu_1 = \mu_2$, are rejected in Kantu and Hokushin, while the nulls that the constants are equal between 1968–1980 and 1981–2000, that is $\mu_2 = \mu_3$, are rejected in Hokkaido, Tohoku, Kantu, Hokushin, Tokai, and Kinki. For the panel as a whole, both nulls are conclusively rejected, and there are significant breaks in the rent-price relationship in both 1967 and 1980. With respect to the first period, 1955–1967, the break in 1967 resulted in an increase of 0.092 (2.573–2.481) of the stationary difference ($r_t - 0.508p_t$), or to a rise of 9% in the rent/price ratio.[11] With respect to the second period, 1968–1980, the break in

[10] Parameter estimates for the panel are averages of the parameters for each region, while the associated t-statistics are calculated by multiplying the sum of t-statistics for each region by $N^{-1/2}$ following Pedroni (2001). DOLS estimates are similar.

[11] Strictly, this is a rise in the ratio $r_t/p_t^{0.51}$.

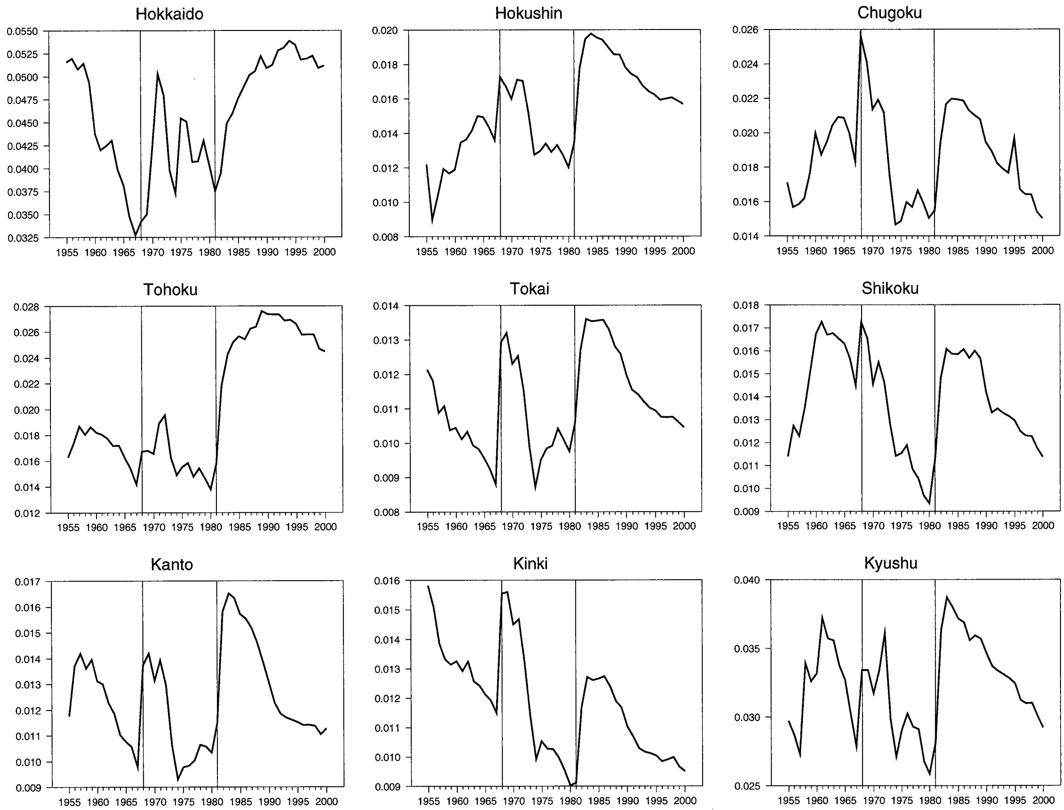

FIGURE 2
RENT/PRICE RATIO, 1955–2000

1980 resulted in a further increase of 0.145 (2.718–2.573) of the stationary difference ($r_t - 0.508p_t$), or to an increase of 15% in the rent/price ratio. Figure 2 intuitively supports these findings. Thus, changes to the land market in 1967 and 1980 both increased rents relative to prices, and there is evidence that Japanese farmland policy has traditionally acted to maintain rents at levels below those implied by free markets.

The rent-price elasticities, β_i, in Table 2 are all positive except for Kantu, which is insignificant. The positive elasticities range from 0.09 for Kinki to 1.29 for Kyushu, but most are in the range 0.39 to 0.70; those for Hokkaido, Tohoku, and Kyushu are also significant. The magnitudes of these elasticities are quite different from each other, and market efficiency differs across regions. The between-group panel FMOLS estimate of the rent-price elasticity is 0.51 and significant. Table 3 shows the results of testing the nulls that $\beta_i = 1$ for each region, which implies that the land market is

TABLE 1
COINTEGRATION TESTS

	Z-LM-Statistic
Hokkaido	0.55 (0.68)
Tohoku	0.73 (0.76)
Kantu	0.55 (0.52)[a]
Hokushin	0.35 (0.77)
Tokai	0.77 (0.73)[a]
Kinki	0.64 (0.65)
Chugoku	0.30 (0.82)
Shikoku	0.32 (0.75)
Kyushu	0.53 (0.77)
Panel	2.12 (2.84)

Note: Bootstrapped critical values at the 5% significance level are in parentheses.
[a] Rejection of the null of cointegration.

TABLE 2
FMOLS Estimates

	β_i	μ_1	μ_2	μ_3
Hokkaido	0.673 (4.86)*	0.826 (0.49)	0.926 (0.53)	1.112 (0.64)
Tohoku	0.699 (4.92)*	−0.076 (0.04)	−0.054 (0.03)	0.399 (0.21)
Kantu	−0.074 (0.27)	9.885 (2.66)*	10.253 (2.68)*	10.717 (2.74)*
Hokushin	0.461 (1.73)	2.859 (0.80)	3.046 (0.85)	3.336 (0.91)
Tokai	0.385 (1.20)	3.668 (0.84)	3.841 (0.87)	4.076 (0.91)
Kinki	0.094 (0.32)	7.801 (1.98)*	7.901 (1.97)*	8.106 (1.98)*
Chugoku	0.463 (1.61)	3.051 (0.81)	3.122 (0.82)	3.224 (0.83)
Shikoku	0.580 (1.62)	1.450 (0.30)	1.370 (0.28)	1.532 (0.31)
Kyushu	1.291 (5.90)*	−7.135 (2.56)*	−7.251 (2.56)*	−7.240 (2.52)*
Panel	0.508 (7.30)*	2.481 (1.76)	2.573 (1.80)	2.718 (2.00)*

Note: Estimated equation [16]: $r_{it} = \mu_1 + \mu_2 + \mu_3 + \beta_i p_{it} + \varepsilon_{it}$ $i = 1,\ldots,N$ $t = 1,\ldots,T$; t-ratios are given in parentheses where $t \sim N(0,1)$.
* Significance at the 5% level.

efficient, against the alternative that $\beta_i \neq 1$. The null is rejected in seven cases and also for the panel, and there is little evidence that the Japanese farmland market is efficient.

Table 4 shows the results of testing for noncausality. We first test the nulls that prices do not cause rents in each region, using the ECMs in [9]. The null H_0^1 in [11] that there is no short- or long-run causality is rejected for Hokkaido, Tohoku, Hokushin, Chugoku, Shikoku, and Kyushu; and the null H_0^2 in [12] that there is no long-run causality is rejected for the same six regions and for Kantu. For the panel, the nulls H_0^3 in [13] that there is no short- or long-run causality and H_0^4 in [14] that there is no long-run causality are both conclusively rejected. Second, we test nulls that rents do not cause prices, using the ECMs in [10]. The nulls H_0^1 of no short- or long-run causality and H_0^2 of no long-run causality are not rejected for all regions. Similarly for the panel, the nulls H_0^3 of no short- or long-run causality and H_0^4 of no long-run causality support the conclusion that rents do not cause prices. We conclude therefore that there is unidirectional causality from prices to rents.

Group mean $\bar{t}(\alpha)$-statistics, which test the null of H_0^5 in [15], also show pervasive long-run causality from prices to rents on average, and homogeneity W-statistics imply that this conclusion is not due to heterogeneous adjustment coefficients, $\alpha_i^1 = \alpha^1$ for $i = 1,\ldots,N$ in [9], that cancel each other out. The battery of causality tests

TABLE 3
Hypothesis Tests

	Wald Statistic		t-Statistic
	$H_0{:}\mu_1 = \mu_2$	$H_0{:}\mu_2 = \mu_3$	$H_0{:}\beta_i = 1$
Hokkaido	1.501 (0.23)	11.365 (0.00)*	2.36 (0.01)*
Tohoku	0.100 (0.75)	89.782 (0.00)*	2.12 (0.02)*
Kantu	8.353 (0.01)*	21.314 (0.00)*	3.92 (0.00)*
Hokushin	7.033 (0.01)*	8.870 (0.00)*	2.02 (0.02)*
Tokai	3.872 (0.06)	8.076 (0.01)*	1.92 (0.03)*
Kinki	1.108 (0.30)	4.088 (0.05)*	3.08 (0.00)*
Chugoku	0.602 (0.44)	1.331 (0.25)	1.87 (0.03)*
Shikoku	0.463 (0.50)	1.720 (0.20)	1.17 (0.12)
Kyushu	2.870 (0.10)	0.031 (0.86)	1.33 (0.09)
Panel	12.708 (0.00)*	83.492 (0.00)*	7.07 (0.00)*

Note: Wald tests are distributed as $F_{1,41}$; t-statistics are distributed as standard normal; panel Wald tests are distributed as χ_1^2; p-values are in parentheses.
* Significance at the 5% level.

TABLE 4
CAUSALITY TESTS

	ECM in [9] H_0: Prices Do Not Cause Rents			ECM in [10] H_0: Rents Do Not Cause Prices		
		SR + LR	LR		SR + LR	LR
	Lags	*F*-Stat.	*t*-Stat.	Lags	*F*-Stat.	*t*-Stat.
Individual Tests						
Hokkaido	3	13.76* (0.00)	−3.51* (0.00)	1	0.87 (0.43)	0.82 (0.41)
Tohoku	3	16.89* (0.00)	−7.09* (0.00)	1	0.78 (0.47)	−1.12 (0.27)
Kantu	1	2.07 (0.14)	−2.02* (0.05)	1	0.10 (0.90)	−0.41 (0.68)
Hokushin	3	13.44* (0.00)	−6.29* (0.00)	1	1.31 (0.28)	−1.56 (0.13)
Tokai	1	0.46 (0.63)	−0.95 (0.35)	1	0.01 (0.99)	0.01 (0.98)
Kinki	1	1.96 (0.15)	−1.47 (0.15)	1	2.61 (0.09)	−0.49 (0.62)
Chugoku	1	5.44* (0.01)	−2.83* (0.00)	1	0.08 (0.92)	−0.31 (0.75)
Shikoku	3	10.68* (0.00)	−4.79* (0.00)	1	1.72 (0.19)	−1.62 (0.11)
Kyushu	1	25.88* (0.00)	−6.95* (0.00)	1	1.56 (0.22)	−1.77 (0.08)
Panel Tests						
LLR-statistic		193.94* (0.00)	136.31* (0.00)		19.84 (0.34)	12.91 (0.16)
Group mean $\bar{t}(\alpha)$-statistic			−3.99* (0.00)			−0.72 (0.24)
Homogeneity *W*-statistic			30.56* (0.00)			5.28 (0.81)

Note: ECM, error-correction model; LR, long run; SR, short run. *p*-Values are in parentheses.
* Significance at the 5% level.

provides overwhelming evidence that prices cause rents, which supports the institutional rent-formation hypothesis, and normalizing the cointegrated relationship in [16] on rents is justified. Our FMOLS estimate of the rent-price elasticity implies that a 1% increase in prices results in a 0.51% increase in rents.

The conclusions that there is cointegration between rents and prices and that there is unidirectional causality from prices to rents is based on estimating the institutional model in [4] and [16]. We now examine the PVM counterpart to [16] where price is hypothesized to be determined by rent. Testing the null of panel cointegration reveals that Z-LM = 2.33 (bootstrapped critical value at the 5% significance level: 3.64), the null is not rejected, and there is cointegration between rents and prices. Testing the null that rents do not cause prices in both the short and long run, the log-likelihood ratio (LLR) = 33.80 (*p*-value: 0.05) and the null is not rejected. Similarly, in the long run, LLR = 15.24 (0.08), the group mean test yields $\bar{t}(\alpha) = -1.13$ (0.13), and the homogeneity test yields $W = 22.12$ (0.01). Thus there is no long-run causality from rents to prices, although the conclusion based on the group mean test may be due to heterogeneous error-correction terms that cancel each other out. Finally, we test the null that prices do not cause rents: in both the short and long run, LLR = 160.28 (0.00); in the long run, LLR = 119.50 (0.00), $\bar{t}(\alpha) = 2.09$ (0.02), and $W = 111.73$ (0.00), and there is clear causality from prices to rents. These results support the conclusions derived from the institutional model.

V. CONCLUSIONS

The farmland market in Japan exhibits some distinctive characteristics. Traditionally, it has been strongly regulated, with rents being determined by a process of institutional governance guided by prices. This contrasts with the framework used in many land market studies, particularly in Europe and North America, where the PVM hypothesizes that land prices are determined by farm rents. And while deregulation and decentralization are under way in the Japanese farmland market, there is still a belief by some that it is inefficient and distorted.

This paper examines the relationship between farmland rents and prices in nine regions in Japan, using annual data for 1955–2000 and recently developed panel cointegration methods that permit structural breaks. Our focus is similar to that of Shigeto, Hubbard, and Dawson (2008) in that we contrast the present-value and institutional rent-formation models by testing the causality between rents and prices, examine the impact on the rent-price nexus of the rent revisions in 1967 and 1980, and test for market efficiency. Whereas Shigeto, Hubbard, and Dawson rely on national time series data of moderate length and conventional cointegration tests, which tend to underreject when structural breaks are present and have poor size and power properties, we seek more robust results and use a richer, panel dataset of regional rents and prices to gain estimation efficiency.

Our results provide evidence of a cointegrating relationship between rents and prices. There are significant breaks in the rent-price relationship following rent revisions and partial deregulation of the farmland market, the effect of which is to increase the long-run equilibrium rent/price ratio by 9% in 1967 and by 15% in 1980. To Shigeto, Hubbard, and Dawson (2008), the break in 1967 is insignificant and that in 1980 led to a 21% increase in the rent/price ratio. We find overwhelming evidence that prices cause rents, which supports the institutional rent-formation hypothesis of Shigeto, Hubbard, and Dawson and runs counter to the PVM.[12] We also find that the rent-price elasticity is 0.51, which is significantly different from unity. Thus, the farmland market in Japan is not efficient, which contrasts with the findings of Shigeto, Hubbard, and Dawson. This result offers some support to those commentators who believe that the Japanese farmland market is inefficient and distorted, owing to government policy and the distinctive nature of the Japanese situation.

[12] A caveat is that we hypothesize either that prices are caused by rents or that rents are caused by prices; we do not examine whether causality changes after rent reform.

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