Newcastle University ePrints

Dawson PJ. [Market Failure and Japanese Farmland Rents](#). *Journal of Agricultural Economics* 2014, 65(2), 406-419.

Copyright:

©2014 The Agricultural Economics Society

This is the peer reviewed version of the following article: Dawson PJ. [Market Failure and Japanese Farmland Rents](#). *Journal of Agricultural Economics* 2014, 65(2), 406-419., which has been published in final form at <http://dx.doi.org/10.1111/1477-9552.12048>. This article may be used for non-commercial purposes in accordance with Wiley Terms and Conditions for self-archiving

Further information on publisher website: www.elsevier.com

Date deposited: 08-10-2014

Version of file: Accepted Author Manuscript

This work is licensed under a [Creative Commons Attribution-NonCommercial 3.0 Unported License](http://creativecommons.org/licenses/by-nc/3.0/)

ePrints – Newcastle University ePrints

<http://eprint.ncl.ac.uk>

Market Failure and Japanese Farmland Rents

P.J. Dawson¹

Abstract

Since the early-1950s, Japanese farmland rents have been regulated and a consensus emerged that rent control led to market failure. Hypothesising a rent-formation model where rents are determined by prices, this paper estimates a threshold autoregressive model which integrates three tests of market failure, namely, inefficiency, bias and asymmetry. There are four results. First, a long-run relationship exists between rents and prices, and the Japanese farmland rental market is efficient. Second, the rent-price elasticity is unity and the market is unbiased. Third, rents are Granger-caused by prices which supports the rent-formation model. Fourth, asymmetry exists where more rapid error-correction occurs immediately after policy reform when rent growth exceeds price growth by 3.6% or more, and rent control has benefitted tenants.

Keywords: Farm rents and land prices, rent control, Japan, market failure, threshold autoregression, efficiency, asymmetric price adjustment, price/rent bias

JEL Classifications: Q15, C32

1. Introduction

The Japanese farmland rental market was reformed after the Second World War as part of the transformation from feudalism to democracy. Three important aspects of this reform were that rents were reduced, rent-in-kind was transformed into money rent, and 'owner-cultivators' were created (Koppel and Kim, 1993). The main instrument of reform was

¹ Phil Dawson is Reader in the School of Agriculture, Food and Rural Development, Newcastle University, Newcastle upon Tyne, NE1 7RU, UK (tel: +44 (0) 0191 222 6883; fax: +44 (0) 0191 222 6720; e-mail: P.J.Dawson@ncl.ac.uk). I am grateful to the Editor and two anonymous referees for the comments on an earlier draft.

the Agricultural Land Law (1952) which *inter alia* had power to control maximum rents. Rent control was revised in 1967 and abolished in 1970 although it was still applied to existing agreements until 1980. Since 1970, a more limited form of rent control has been exercised by local Chambers of Agriculture which oversee rents and exert influence when they are regarded as too high (Sekiya, 2002). By contrast, farmland prices have not been directly controlled although agricultural land is not allowed to convert to other uses to prevent speculation. Honma (1994), Egaitsu and Shogenji (1995) and Kusakari (1998) argue that rent control protects tenants against rent rises; the Japanese farmland rental market therefore is inefficient and there is market failure.² A counterview is that rent control restricts short-run attempts by landowners to raise rents above the efficient, long-run equilibrium level, and rent control may have improved market efficiency rather than exacerbated it. Whether the Japanese farmland rental market is inefficient therefore is unclear.

Empirical evidence on the existence of market failure is mixed. Shigeto *et al.* (2008) use national data for 1955-2000 and standard cointegration methods to estimate a rent-formation model where rent is set as a simple mark-down on price through a process of institutional governance.³ They focus on two aspects of market failure and test for market

² The market is further distorted: before 1970, terminating tenancy contracts was difficult because permission of the governor of the prefecture was required; and since 1970, landowners could only terminate contracts that had lasted at least 10 years.

³ Most of the empirical literature on the farmland rent-price relationship elsewhere is based on the present valuation model whereby rent determines price. The model hypothesises that price is equal to the capitalised value of future income streams or rent, that is, $P_t = \alpha \sum_{j=1}^{\infty} \alpha^j E_t[R_{t+j}]$ where P_t is the price at the beginning of time period t , R_t is 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 expectations operator based on information available at time t . A constant expected discount rate is commonly assumed which implies that there is a long-run relationship between price and rent, that is, $p = \beta_0 + \beta_1 r$ where $p = \ln P^*$, $r = \ln R^*$, $\ln$ is the natural logarithm, P^* and R^* are the long-run equilibrium price and a constant expectation of

efficiency when rents and prices are cointegrated, and unbiasedness when the rent-price elasticity is unity (Kellard, 2002).⁴ There are four results. First, rents and prices are cointegrated and the Japanese farmland rental market is efficient. Second, prices Granger-cause rents which supports the rent-formation model. Third, the rent-price elasticity is unity and the market is unbiased. Fourth, there is a structural break in the rent-price relationship in 1980 when the rent/price ratio rose as maximum rents were finally abolished. By contrast, Sanjuán *et al.* (2009) use panel cointegration methods and regional data for 1955-2000. Their results show evidence of cointegrating relationships in each region with structural breaks in both 1967 and 1980, prices Granger-cause rents which again supports the rent-formation model, and rents are inelastic with respect to price which implies market bias and rent control favours tenants.

The rent-formation models estimated by Shigeto *et al.* (2008) and Sanjuán *et al.* (2009) hypothesise symmetric price transmission where a rent rise following a price increase is the same absolute magnitude as a rent fall following a corresponding price decrease, and where the speed of adjustment to long-run equilibrium is the same in each case. A third dimension of market failure is asymmetric adjustment of rents to prices and there are two reasons why asymmetry may exist. First and drawing on Currie (1981, pp.87-99), asymmetry may arise because of land fixity, the relative bargaining strengths of landowners and tenants and asymmetric information, and large landowners with many tenants are able to exploit local monopoly power. Rents therefore are 'sticky' in a downwards direction with rents rising faster than they fall. By contrast, rents maybe 'sticky' in an upwards direction when long leases exist

equilibrium rent, $\beta_1=1$ and $\beta_0=\ln(1/i)$. Empirical applications include Falk (1991), Lloyd *et al.* (1991), Lloyd (1994) and Lence and Miller (1999).

⁴ Shigeto *et al.* (2008) use a more general definition of market efficiency where both rents and prices are cointegrated and the rent-price elasticity is unity.

and rent rises are slower than rent falls: tenant farmers suffer when there is an unanticipated fall in commodity prices and they seek immediate rent reductions; conversely, landowners suffer when there is an unanticipated rise in commodity prices since they have to wait for contracts to expire before rents can be increased. A second reason for asymmetry is the efficacy of rent control before 1980 under the Agricultural Land Law and thereafter from the activities of the Chambers of Agriculture. Any difference between this efficacy and the speed at which landowners can increase short-run rents above their efficient, long-run equilibrium level can lead to asymmetric rent adjustment.

This paper re-examines the farmland rent-price relationship in Japan using the rent-formation model of Shigeto *et al.* (2008). By contrast to previous studies, we test for three aspects of market failure, namely inefficiency, bias and asymmetry, within the integrated empirical framework of Enders and Siklos (2001). Evidence of market failure has implications about the effects of rent control. If the farmland rental market is inefficient, there is no long-run relationship between rents and prices, and price does not determine rent. If the market is biased, rent changes more or less than proportionately in response to a price change, and rent control favours either landowners or tenants. Third, asymmetry implies that adjustment when rents are too high is either more or less rapid than adjustment when rents are too low. If adjustment is more rapid when rents are too high than when they are too low, rent control favours tenants; otherwise, landowners are favoured. The paper is structured as follows: Section 2 provides some background on incorporating asymmetry into a cointegration framework which is a key focus, Section 3 discusses the empirical model and method, Section 4 reports the results, and Section 5 concludes.

2. Asymmetric Price Transmission: A Selected Review

There is a growing literature on asymmetric price transmission although there appears to be none on farmland markets. Much is concerned with agricultural markets although other lines of inquiry include pricing behaviour in fuel/energy and financial markets. Meyer and von Cramon-Taubadel (2004) and Frey and Manera (2007) review econometric models and the former also survey some applications, particularly in agriculture. It is not the intention to replicate these reviews, but rather to provide some background to the case at hand.

Price transmission examines the relationship between prices either at different stages of the marketing chain or between spatially separated markets. Consider a linear relationship between two prices, p_1 and p_2 :

$$p_1 = \beta_0 + \beta_1 p_2 \tag{1}$$

where β_0 and β_1 are parameters and p_1 is a mark-up on p_2 . Asymmetric price response is concerned with either the magnitude of price transmission and/or the speed of adjustment (Meyer and von Cramon-Taubadel, 2004). Magnitude symmetry is where β_1 is the same value for an increase in p_2 as for a corresponding fall, whereas magnitude asymmetry is where β_1 differs between these two cases. Asymmetric adjustment on the other hand is where p_1 reacts more rapidly to an increase in p_2 than to a fall, or *vice versa*. If p_1 rises faster when p_2 increases than when p_2 falls, p_1 is 'sticky' in a downwards direction and this is 'positive asymmetric price transmission' (Peltzman, 2000) which is referred to in the informal literature as 'rockets and feathers'.

Early work on asymmetric relationships, particularly by Tweeten and Quance (1969), Wolfram (1971) and Houck (1977), examines magnitude asymmetry in agricultural supply response while later studies including Boyd and Brorsen (1988), Bailey and Brorsen (1989) and Mohanty *et al.* (1995) consider magnitude asymmetric price transmission. The notion of magnitude asymmetry however is criticised by von Cramon-Taubadel and Loy (1996) and by von Cramon-Taubadel (1998) because it may produce spurious regression results or is incompatible with cointegration. Specifically, no unique long-run relationship can exist because magnitude asymmetry implies a permanent difference between positive and negative periods of transmission, and there are two long-run relationships, one when p_2 is rising and the other when p_2 is falling. Accordingly, more recent literature focuses on asymmetric adjustment towards long-run equilibrium within a cointegration framework. Building on Tsay's (1989) threshold autoregressive (TAR) model, threshold cointegration models allow the speed of adjustment to differ depending on whether the deviation from long-run equilibrium is above or below a threshold. Examples include Balke and Fomby (1997), Engle and Granger (1998), Enders and Siklos (2001), Hansen and Seo (2003), and Greb *et al.* (2013).

Threshold cointegration methods have been used to examine price transmission at different stages of the marketing food chain. For example, Goodwin and Holt (1999) and Goodwin and Harper (2000) examine US wholesale-retail prices relationships for beef and pork respectively and both find that shocks are unidirectional from wholesale to retail prices. In addition, the former find modest asymmetry while the latter find varying asymmetric strengths. More evidence of asymmetry is found by Cramon-Taubadel (1998) between German producer and wholesale pork prices, by Abdulai (2002) between Swiss producer and retail pork prices, by Alemu and Ogundeji (2010) between South African producer and retail

food prices, and by Falkowski (2010) between Polish producer and retail liquid milk prices. In each case, an increase in the producer price that squeezes the margin is transmitted faster to the wholesale/retail price than is a decrease in the producer price that increases the margin.

Evidence of asymmetric price transmission in spatially separated agricultural markets is mixed. For example, Abdulai (2000) finds that wholesale maize prices in local markets in Ghana respond faster to increases in the central market price than to decreases. Subervie (2011) estimates relationships between producer coffee prices in El Salvador, India and Colombia and world prices and finds that positive asymmetry is replaced by negative asymmetry conditional on domestic liberalisation reforms. By contrast, little evidence of asymmetry is found by von Cramon-Taubadel and Loy (1996) in bivariate relationships between the US wheat price and corresponding prices in Argentina, Australia, Canada and the EU, and by Goodwin and Piggott (2001) for US corn and soybean prices in local and central markets.

Other econometric methods have also been used to examine asymmetric price transmission and examples include Serra and Goodwin (2003) and Awokuse and Wang (2009). The former apply threshold vector error-correction models and find asymmetric bivariate relationships between farm and retail prices for various dairy products in Spain. The latter apply threshold TAR cointegration models to examine producer-retail price relationships for US butter, cheese, and liquid milk, and there is strong evidence of asymmetry for butter and liquid milk but not for cheese.

Overall, there is much evidence of both symmetric and asymmetric price transmission. Meyer and von Cramon-Taubadel (2004) observe almost equal conclusions in

205 tests reviewed, but of the 10 tests that use threshold cointegration models, eight reject symmetry which suggests that asymmetry is perhaps more easily identified by such methods.

3. Empirical Model and Method

Adapting (1), Shigeto *et al.* (2008) propose a simple rent-formation model where farmland rent is set in accordance with land price:

$$r_t = \beta_0 + \beta_1 p_t + u_t \quad t=1, \dots, n \quad (2)$$

where r_t is the rent paid in period t , p_t is the price at the beginning of time period t , r_t and p_t are defined in natural logarithms, and u_t is an error term. *A priori*, we expect that $\beta_0 < 0$ and $\beta_1 > 0$. Define market efficiency where r_t and p_t are cointegrated, while rents are unbiased if $\beta_1 = 1$ where a 1% increase in price leads to a 1% increase in rent.

Standard econometric models of cointegrated variables are those of Engle and Granger (1987) and Johansen (1988).⁵ These methods assume that adjustment is symmetric, that is, the speed of adjustment from disequilibrium is linear, and positive and negative errors induce the same speed of adjustment back to long-run equilibrium. Consider the Engle-Granger method. Conditional on r_t and p_t being non-stationary I(1) variables, the long-run relationship in (2) is estimated by ordinary least squares (OLS). If u_t (which may be serially correlated) is stationary, that is I(0), then r_t and p_t are cointegrated and a long-run relationship

⁵ Shigeto *et al.* (2008) use Johansen *et al.*'s (2000) method which extends Johansen (1988) to permit structural breaks.

exists between them. The null of no cointegration between r_t and p_t is tested by testing the stationarity of u_t using:

$$\Delta u_t = \rho_1 u_{t-1} + \sum_{i=1}^{s-1} \gamma_i \Delta u_{t-i} + v_t \quad (3)$$

where u_t are the errors in (2) and v_t is a white noise error term. Lagged values of Δu_t may be added to (3) to ensure that v_t approximates white noise. The null of a unit root and thus of no cointegration is $\rho_1=0$. If $\rho_1<0$, u_t is stationary and r_t and p_t are cointegrated and the Granger representation theorem implies that a corresponding error-correction model (ECM) exists (Engle and Granger, 1987).

Engle-Granger's test of no cointegration is mis-specified if adjustment is not symmetric and Enders and Siklos (2001) extend the model to admit asymmetric threshold adjustment. They develop two cointegration models, namely a threshold autoregressive (TAR) model and a momentum threshold autoregressive (M-TAR) model. In the TAR model, (3) is replaced by:

$$\Delta u_t = \rho_1 I_t u_{t-1} + \rho_2 (1 - I_t) u_{t-1} + \sum_{i=1}^{s-1} \gamma_i \Delta u_{t-i} + v_t \quad (4)$$

where I_t is the Heaviside indicator which depends on the level of u_{t-1} :

$$I_t = \begin{cases} 1 & \text{if } u_{t-1} \geq \tau \\ 0 & \text{if } u_{t-1} < \tau \end{cases} \quad (5)$$

and τ is the value of the threshold. It is sometimes reasonable to set $\tau=0$ in (5) so that adjustment is $\rho_1 u_{t-1}$ if u_{t-1} is above its long-run equilibrium value and is $\rho_2 u_{t-1}$ if u_{t-1} is below its long-run equilibrium value. Thus, u_t is permitted to exhibit more momentum in one direction than the other. For example, if $|\rho_1| > |\rho_2|$, adjustment from positive values of u_{t-1} is more rapid than adjustment from corresponding negative values. Alternatively if $\rho_1 = \rho_2$, the Engle-Granger model in (2) and (3) is nested in the TAR model of (2), (4) and (5). The M-TAR model is similar to the TAR model except that the Heaviside indicator depends upon the change in u_{t-1} , and (5) is replaced by:

$$I_t = \begin{cases} 1 & \text{if } \Delta u_{t-1} \geq \tau \\ 0 & \text{if } \Delta u_{t-1} < \tau \end{cases} \quad (6)$$

Here, adjustment is $\rho_1 u_{t-1}$ if $\Delta u_{t-1} \geq \tau$, and is $\rho_2 u_{t-1}$ if $\Delta u_{t-1} < \tau$. Again, it may be reasonable to set $\tau=0$ in (6) but the more general case is where τ is unconstrained and it needs to be estimated. Enders and Siklos recommend a search over possible thresholds lying in the middle 70% of the range of Δu_t , and τ is chosen to minimise the residual sum of squares.

In both TAR and M-TAR models, necessary and sufficient conditions for the stationarity of u_t in (4) when r_t and p_t are cointegrated are that $\rho_1 < 0$, $\rho_2 < 0$ and $(1+\rho_1)(1+\rho_2) < 1$ for any value of τ , which imply that $u_t=0$ is the long-run equilibrium value of the system. If r_t and p_t are not cointegrated, there is no threshold, τ , and $\rho_1=0$ and/or $\rho_2=0$. Enders and Siklos propose two tests of the null of no cointegration where the alternative is cointegration with asymmetry where $u_t \sim I(0)$, and $\rho_1 < 0$ and $\rho_2 < 0$ in (4). The first is the t-Max test. Denoting the larger of the t-statistics associated with ρ_1 and ρ_2 as t-Max, the null is either that $\rho_1=0$ or that

$\rho_2=0$. Second, the Φ -statistic tests the joint null that $\rho_1=\rho_2=0$. Both tests have non-standard distributions and Enders and Siklos provide critical values from Monte Carlo experiments. They also argue that the Φ -statistic has more power and is preferred.

The corresponding short-run ECM for rents in both TAR and M-TAR models is:

$$\Delta r_t = \theta + \alpha_1 I_t u_{t-1} + \alpha_2 (1 - I_t) u_{t-1} + \sum_{i=1}^{k-1} \delta_{1i} \Delta r_{t-i} + \sum_{i=1}^{k-1} \delta_{2i} \Delta p_{t-i} + \varepsilon_t. \quad (7)$$

Lagged values of Δr_t and Δp_t are added to (7) to ensure that the error term, ε_t , approximates white noise. In (7), α_1 is the speed of adjustment when disequilibrium is above the threshold, and for the TAR (M-TAR) model this is when $u_{t-1} \geq \tau$ ($\Delta u_{t-1} \geq \tau$) when long-run rents are too high. Similarly, α_2 is the speed of adjustment coefficient when disequilibrium is below the threshold, and for the TAR (M-TAR) model this is when $u_{t-1} < \tau$ ($\Delta u_{t-1} < \tau$) when long-run rents are too low. Positive discrepancies from long-run equilibrium are shorter lived in the M-TAR model.

4. Data and Results

National average farmland rents and prices from actual transactions are not available. Instead, annual surveys of rents and prices are carried-out by the Japan Real Estate Institute on a particular census day (31 March) whereby questionnaires are administered to agricultural experts employed at municipal halls or belonging to agricultural councils including Chambers of Agriculture. Rents are from farmland leases, and prices are for owner-cultivated farmland from sales. One-season farmland leases, orchards, tea plantations, and mulberry fields are excluded. The sample period is 1955-2010 (56 observations) and these

data are the most consistent and longest time series of national farmland rents and prices available. Nominal average rents and prices (yen/10are, where 100are=1hectare) are deflated by the GDP deflator and are for 'good' paddy and vegetable fields. They are shown in Figure 1 and co-movement is apparent..

Figure 1 about here

Augmented Dickey-Fuller tests (Dickey and Fuller, 1981) are used to test for unit roots in r_t and p_t . The number of lags in each equation, with a maximum of four, is chosen using the Akaike Information Criterion (AIC). Test statistics when a trend is included in each equation are -1.36 for r_t and -1.18 for p_t (critical value at the 90% confidence level: -3.15) while corresponding statistics from non-trended equations are -1.78 and -1.39 (critical value: -2.58). The null of a unit root in each case is not rejected irrespective of whether a trend is included or not and both r_t and p_t are non-stationary I(1) variables.

The long-run relationship in (2) is common to the Engle-Granger, TAR and M-TAR models and the OLS estimate is:

$$r_t = -6.694 + 1.186p_t + \hat{u}_t \quad R^2=0.76, D.W.=0.29 \quad (8)$$

The residuals, $\hat{u}_t$, from (8) are used to test the null of no cointegration in all three models and the results are shown in Table 1. In each case, the AIC indicates that three lags are appropriate. The Engle-Granger cointegration test statistic is $t=-2.43$ (critical value at 90% confidence interval: c.v.=-2.45 (MacKinnon, 1994)), and the null of no cointegration between r_t and p_t is not rejected. In the TAR model in (4) and (5), we test the null of no cointegration

with $\tau=0$. Point estimates of ρ_1 and ρ_2 imply that there is convergence of rents towards long-run equilibrium after a shock to price, but $\Phi=2.89$ (c.v.=5.22) and t-Max=-1.84 (c.v.=-1.89) and the null is not rejected. Next, we test the null of no cointegration in the M-TAR model in (4) and (6) with $\tau=0$. Point estimates of ρ_1 and ρ_2 again imply convergence. The conclusions from the Φ -statistic and the t-Max test are mutually reinforcing: $\Phi=4.08$ (c.v.=5.32) while t-Max=-1.83 (c.v.=-1.84) and null is again not rejected. Finally, we test the null of no cointegration in the M-TAR model in (4) and (6) without restricting τ . A search grid yields $\tau=0.036$, and point estimates of ρ_1 and ρ_2 imply convergence. The conclusions from the Φ -statistic and the t-Max test are again mutually reinforcing but here, $\Phi=6.13$ (c.v.=5.99) and t-Max=-1.97 (c.v.=-1.72), the null of no cointegration is rejected, and the farmland market is efficient. AICs indicate that the more general M-TAR model with an unrestricted threshold, τ , best fits the data.⁶ In this preferred model, the null that $\rho_1=\rho_2$ is rejected and symmetric adjustment in the Engle-Granger model is not supported.

Diagnostic mis-specification tests are also shown in Table 1. Denote the Breusch-Godfrey test for fourth-order serial correlation by χ^2_{corr} (Greene, 2012, pp.316-317), Engle's test for autoregressive conditional heteroscedasticity of order four by χ^2_{ARCH} (Greene, pp.976-977), White's test for heteroscedasticity by χ^2_{het} (Greene, pp.315-316), and Doornik and Hansen's (2008) normality test by E_p . Tests in all equations suggest generally well-specified models although non-normality is evident and statistical tests need to be treated with caution. The preferred M-TAR model with $\tau \neq 0$ (Table 1, fifth column) implies that a long-run relationship exists between rents and prices and that adjustment to long-run equilibrium is

⁶ This pattern of cointegration between the four models is similar to that of Enders and Siklos who examine bivariate relationships between US interest rates.

asymmetric. Since the null that $\rho_1 = \rho_2$ is rejected and $|\rho_1| > |\rho_2|$, the speed of adjustment to positive discrepancies from long-run equilibrium when $\Delta \hat{u}_{t-1} \geq 0.036$ is more rapid than adjustment to negative discrepancies when $\Delta \hat{u}_{t-1} < 0.036$.

Table 1 about here

Speeds of adjustment of rents to price shocks are contained in the ECM for rent in (7). AICs rise as the number of lags increase to four but both the Schwarz Bayesian and Hannan-Quinn criteria indicate one lag and the estimated ECM for rent is:

$$\Delta r_t = 0.001 - 0.479I_t\hat{u}_{t-1} - 0.154(1 - I_t)\hat{u}_{t-1} + 0.317\Delta r_{t-1} + 0.392\Delta p_{t-1} + \hat{\varepsilon}_t \quad (9)$$

(0.10) (-2.37) (-1.65) (1.92) (1.06)

$$R^2=0.27; F_{4,49}=4.55 [0.00]; D.W.=2.01;$$

$$\chi^2_{corr}=8.12 [0.09]; \chi^2_{ARCH}=0.27 [0.99]; \chi^2_{het}=49.31 [0.24]; E_p=24.82 [0.00]$$

(t-statistics in parentheses; p-values in square brackets)

Diagnostic mis-specification tests suggest that (9) is reasonably well-specified although non-normality is evident. Following Enders (2010, pp.365-371), we test the null of no Granger-causality where both speed of adjustment coefficients and Δp_{t-1} do not enter the Δr_t -equation. The test statistic is $F_{3,49}=4.03$ (p-value: 0.01) and we conclude that r_t is Granger-caused by p_t . In the ECM for Δp_t which is not reported, there is evidence of non-normality but the null of no Granger-causality yields $F_{3,49}=0.75$ (p-value: 0.53) and there is strong evidence that r_t does not Granger-cause p_t . We conclude that p_t is weakly exogenous and there is support for the rent-formation model.

We also test for two structural breaks, in 1967 when the maximum rent for new rental agreements was revised, and in 1980 when maximum rents were abolished. In the preferred M-TAR model, two dummy variables for 1969-1980 and 1981-2010 are included in (4); testing the joint significance of the associated parameters yields $F_{3,49}=0.08$ (p-value: 0.92) and there is no evidence of breaks. Finally, we test the null of unbiasedness, that is $\beta_1=1$ in (8). Following Saikkonen (1991) and Stock and Watson (1993), the long-run relationship is estimated by dynamic ordinary least squares (DOLS) where (8) is augmented by lags and leads. Admitting autocorrelation up to order two and lags and leads up to three, AICs suggest that correcting for first-order autocorrelation and two lags/leads is appropriate. The Wald statistic is $\chi_1^2=0.17$ (p-value: 0.68) and the null is not rejected. Thus, the Japanese farmland rental market is unbiased and a 1% increase in the price leads to a 1% increase in rents in the long run.⁷

Speeds of adjustment of rents towards long-run equilibrium following price shocks are indicated by the estimates of α_1 and α_2 in (9) and are illustrated in Figure 2. For positive discrepancies from long-run equilibrium when $\Delta\hat{u}_{t-1}\geq 0.036$, 48% of the adjustment to long-run equilibrium takes place in the first year, 93% of full adjustment is complete by the end of fourth year, and full adjustment is complete by the end of ninth year. For negative discrepancies from long-run equilibrium when $\Delta\hat{u}_{t-1}<0.036$, only 15% of the adjustment to long-run equilibrium takes place in the first year, 49% of full adjustment is complete by the

⁷ The specification of the rent-formation model in (2) is parsimonious and asymmetry may be the result of omitted variables, particularly the real interest rate. To test this hypothesis, (2) is augmented by the real interest rate and the significance of its associated parameter is tested using DOLS with data for 1961-2010. Correcting for second-order autocorrelation with one lag/lead, the Wald statistic is $\chi_4^2=0.06$ (p-value: 0.81), and there is no evidence that the real interest rate determines rent. Thus, the simple rent-formation model in (2) appears reasonable.

end of the fourth year, and full adjustment takes 32 years.⁸ Thus, the speed of adjustment resulting from positive discrepancies from long-run equilibrium is more rapid than adjustment from corresponding negative discrepancies. Figure 3 shows the estimated changes in the residuals, $\Delta\hat{u}_t$, and $\tau=0.036$, and more rapid adjustment when $\Delta\hat{u}_{t-1}\geq 0.036$ occurs in 1969, 1972 and 1982-84 which follows the policy reforms in 1967, 1970 and 1980, while adjustment is slower when policy is unchanged when $\Delta\hat{u}_{t-1}<0.036$. The findings of efficiency and unbiasedness, where $\beta_1=1$, and the preference for the M-TAR model imply that the adjustment of rents to long-run equilibrium from above (when rents are too high and $\Delta\hat{u}_{t-1}\geq 0.036$) occurs when $\Delta r_{t-1}-\Delta p_{t-1}\geq 0.036$. Conversely, if rents are too low and $\Delta\hat{u}_{t-1}<0.036$, then $\Delta r_{t-1}-\Delta p_{t-1}<0.036$. Thus, more rapid adjustment occurs when rent growth exceeds price growth by 3.6% or more.

Figures 2 and 3 about here

5. Conclusions

This paper examines the farmland rent-price relationship in Japan using national data for 1955-2010 and Shigeto *et al.*'s (2008) rent-formation model. The aim is to examine market failure, and in particular to test for inefficiency, bias and asymmetry. There are four results. First, rents and prices are cointegrated, a long-run relationship exists between them and the Japanese farmland rental market is efficient. Second, rents are Granger-caused by prices. These two findings confirm those of both Shigeto *et al.* (2008) and Sanjuán *et al.*

⁸ A *caveat* is that adjustment coefficients have poor small-sample properties (Hansen, 1997; and Enders *et al.* (2007)).

(2009) and they support the rent-formation model.⁹ Third, the Japanese farmland rental market is unbiased and a 1% increase in the farmland price leads to a 1% increase in rents. This concurs with Shigeto *et al.* but contrasts with Sanjuán *et al.* who find an inelastic rent-price relationship and market bias. These different findings arise from different data, different samples and different empirical methods.

A third dimension of market failure not addressed elsewhere is asymmetry and our fourth result is that there is evidence of asymmetric adjustment of rents to prices where a momentum threshold autoregressive model best describes the data. In particular, there is positive asymmetric transmission where adjustment from positive discrepancies from long-run equilibrium is more rapid than adjustment from negative discrepancies. For the former, 48% of the adjustment to long-run equilibrium takes place each year and full adjustment takes nine years; while for the latter, only 15% of the adjustment takes place each year, and full adjustment takes over 30 years. Moreover, the adjustment of rent to a price shock is more rapid when rent growth exceeds price growth by 3.6% or more. Japanese farmland policy reforms were implemented in 1967, 1970 and 1980, and periods of more rapid adjustment occur two years later than each reform and were short-lived. Following the more substantive reforms, land prices rose by 9% in 1968, and by 10% over the three years 1981-83 and these compare with an average growth rate of 0.8% per annum over the sample period. Uncertainty about the effects of these reforms, combined with land fixity, the relative bargaining strengths of landowners and tenants and asymmetric information, allowed landowners to exploit monopoly power and there were substantial increases in rent of 42% in 1968 and 69% during

⁹ Shigeto *et al.* (2008) and Sanjuán *et al.* (2009) use samples for 1955-2000. The longer sample used here for 1955-2010 do not disturb their conclusion that rents are Granger-caused by prices.

1981-83 which compare with an average annual growth rate of 0.5%. With respect to the current policy regime, the Chambers of Agriculture appeared slow in adjusting to the reform of 1980: as statutory rent control was relaxed, rents increased rapidly but then the Chambers exerted increasing influence and imposed rapid, downward pressure countervailing the monopoly power initially enjoyed by landowners. Since the mid-1980s, rents and prices have been more aligned: they are relatively stable with low (mainly negative) growth rates, rent growth is below price growth, and adjustment has been slower.

In summary, we find no evidence that the Japanese farmland rental market is inefficient or biased. However, market failure is present in the form of asymmetric adjustment of rents to price shocks and rent control is faster at smoothing-out large increases in rents than it is at smoothing-out large decreases. Rent control has been beneficial to tenant farmers and detrimental to landowners.

References

- Abdulai, A. (2000). Spatial price transmission and asymmetry in the Ghanaian maize market, *Journal of Development Economics*, 63(2), 327-349.
- Abdulai, A. (2002). Using threshold cointegration to estimate asymmetric price transmission in the Swiss pork market, *Applied Economics*, 34(6), 679-687.
- Alemu, Z.G. and Ogundeji, A.A. (2010). Price transmission in the South African food market, *Agrekon*, 49(4), 433-445.
- Awokuse, T.O. and Wang, X. (2009). Threshold effects and asymmetric price adjustments in US dairy markets, *Canadian Journal of Agricultural Economics*, 57(2), 269-286.
- Bailey, D. and Brorsen, B.W. (1989). Price asymmetry in spatial fed cattle markets, *Western Journal of Agricultural Economics*, 14(2), 246-252.
- Balke, N. and Fomby, T. (1997). Threshold cointegration, *International Economic Review*, 38(3), 627-645.
- Boyd, S.M. and Brorsen, B.W. (1988). Price asymmetry in US pork marketing channel, *North Central Journal of Agricultural Economics*, 10(1), 103-110.
- Currie, J.M. (1981). *The Economic Theory of Agricultural Land Tenure*. Cambridge: Cambridge University Press.
- Dickey, D.A. and Fuller, W.A. (1981). Likelihood ratio statistics for autoregressive time series with a unit root, *Econometrica*, 49(4), 1057-1072.
- Doornik, J.A. and Hansen, H. (2008). An omnibus test for univariate and multivariate normality, *Oxford Bulletin of Economics and Statistics*, 70(Supplement 1), 927-939.
- Egaitsu, N., and Shogenji, S. (1995). *A New Proposal for Enrichment of Japan's Agricultural Industry*. Tokyo: Ieno Hikari Association (in Japanese).
- Enders, W. (2010). *Applied Econometric Time Series*. New York: John Wiley & Sons.

- Enders, W. and Falk, B.L. (1999). *Confidence Intervals for Threshold Autoregressive Models*. Working paper, Department of Economics, Iowa State University.
- Enders, W. and Siklos, P.L. (2001). Cointegration and threshold adjustment, *Journal of Business and Economic Statistics*, 19(2), 166-176.
- Enders, W., Falk, B.L. and Siklos, P. (2007). A threshold model of real US GDP and the problem of constructing confidence intervals in TAR models, *Studies in Nonlinear Dynamics and Econometrics*, 11, 1-28.
- Engle, R.F. and Granger, C.W.J. (1987). Cointegration and error correction: representation, estimation and testing, *Econometrica*, 55(2), 251-276.
- Engle, R.F. and Granger, C.W.J. (1998). Unit-root tests and asymmetric adjustment with an example using the term structure of interest rates, *Journal of Business and Economic Statistics*, 16(3), 304-311.
- Falk, B. (1991). Formally testing the present value model of farmland prices, *American Journal of Agricultural Economics*, 73(1), 1-10.
- Falkowski, J. (2010). Price transmission and market power in a transition context: evidence from the Polish fluid milk sector, *Post-Communist Economies*, 22(4), 513–529.
- Frey, G. and Manera, M. (2007). Econometric models of asymmetric price transmission, *Journal of Economic Surveys*, 21(2), 349-415.
- Goodwin, B.K. and Holt, T. (1999). Price transmission and asymmetric adjustment in the US beef sector, *American Journal of Agricultural Economics*, 81(3), 630-637.
- Goodwin, B.K. and Harper, D.C. (2000). Price transmission, threshold behavior, and asymmetric adjustment in the US pork sector, *Journal of Agricultural and Applied Economics*, 32(3), 543–553.

- Goodwin, B.K. and Piggott, N.E. (2001). Spatial market integration in the presence of threshold effects, *American Journal of Agricultural Economics*, 83(2), 302-317.
- Greb, F., von Cramon-Taubadel, S., Krivobokova, T. and Munk, A. (2013). The estimation of threshold models in price transmission analysis, *American Journal of Agricultural Economics*, 95(4), 900–916.
- Greene, W.H. (2012). *Econometric Analysis*. Harlow: Pearson Education Limited.
- Hansen B.E. (1997). Inference in TAR models, *Studies in Nonlinear Dynamics and Econometrics*, 2(1), 119-131.
- Hansen B.E. and Seo, B. (2003). Testing for two-regime threshold cointegration in vector error-correction models, *Journal of Econometrics*, 110(2), 293-318.
- Honma, M. (1994). *Political Economy of Agricultural Problems*. Tokyo: Nihon Keizai Shimbun, Inc. (in Japanese).
- Houck, J.P. (1977). An approach to specifying and estimating nonreversible functions, *American Journal of Agricultural Economics*, 59(3), 570-572.
- Japan Real Estate Institute. *The Survey on Land Prices and Farm Rents*. Tokyo: JREI.
- Johansen, S. (1988). Statistical analysis of cointegrating vectors, *Journal of Economic Dynamics*, 12(2-3), 231-254.
- Johansen, S., Mosconi, R. and Nielsen, B. (2000). Cointegration analysis in the presence of structural breaks in the deterministic trend, *Econometrics Journal*, 3(2), 216-249.
- Kellard, N. (2002). Evaluating commodity market efficiency: are cointegration tests appropriate? *Journal of Agricultural Economics*, 53(3), 513–529.
- Koppel, B. and Kim, D.Y. (1993). *Land Policy Problems in East Asia - Toward New Choices - A Comparative Study of Japan, Korea and Taiwan*. Honolulu, HI: East-West Center; and Kyounggi-Do, Republic of Korea: Korea Research Institute for Human Settlements.

- Kusakari, H. (1998). Japanese rice farming and the development of rice policy. In M. Okuno, M. and Honma, M. eds., *Economic Analysis of Agricultural Problems*. Tokyo: Nihon Keizai Shimbun, 115-141 (in Japanese).
- Lence, S.H. and Miller, D.J. (1999). Transactions costs and the present value model of farmland: Iowa, 1900-1994, *American Journal of Agricultural Economics*, 81(2), 257-272.
- Lloyd, T.A. (1994). Testing present value model of agricultural land values, *Oxford Bulletin of Economics and Statistics*, 56(2), 209-223.
- Lloyd, T.A., Rayner, A.J. and Orme, C.D. (1991). Present-value models of land prices in England and Wales, *European Review of Agricultural Economics*, 18(2), 141-166.
- MacKinnon, J.G. (1994). Approximate asymptotic distribution functions for unit-root and cointegration tests, *Journal of Business and Economic Statistics*, 12(2), 167-176.
- Meyer, J. and von Cramon-Taubadel, S. (2004). Asymmetric price transmission: a survey, *Journal of Agricultural Economics*, 55(3), 581-611.
- Mohanty, S., Peterson, E.W.F. and Kruse, N.C. (1995). Price asymmetry in the international wheat market, *Canadian Journal of Agricultural Economics*, 43(3), 355-366.
- Peltzman, S. (2000). Prices rise faster than they fall, *Journal of Political Economy*, 108(3), 466-502.
- Saikkonen, P. (1991). Asymptotically efficient estimation of cointegration regressions, *Econometric Theory*, 7, 1-21.
- Sanjuán, A.I., Dawson, P.J., Hubbard, L.J. and Shigeto, S. (2009). Rents and land prices in Japan: a panel cointegration approach, *Land Economics*, 85(4), 587-597.
- Sekiya, S. (2002). *Japanese Land System*. Tokyo: Japan: Agriculture Policy Research Association (in Japanese).

- Serra, T. and Goodwin, B.K. (2003). Price transmission and asymmetric adjustment in the Spanish dairy sector, *Applied Economics*, 35(18), 1889-1899.
- Shigeto, S., Hubbard, L.J. and Dawson, P.J. (2008). On farmland prices and rents in Japan, *Agricultural Economics*, 39(1), 103-109.
- Stock, J.H. and Watson, M.W. (1993). A simple estimator of cointegrating vectors in higher order integrated systems, *Econometrica*, 61(4), 783-820.
- Subervie, J. (2011). Producer price adjustment to commodity price shocks: an application of threshold cointegration, *Economic Modelling*, 28(5), 2239-2246.
- Tsay, R.S. (1989). Testing and modeling threshold autoregressive processes, *Journal of the American Statistical Association*, 84(405), 1-240.
- Tweeten, L.G. and Quance, C.L. (1969). Positivistic measures of aggregate supply elasticities: some new approaches, *American Journal of Agricultural Economics*, 51(2), 342-352.
- von Cramon-Taubadel S. and Loy, J.P. (1996). Price asymmetry in the international wheat market: comment, *Canadian Journal of Agricultural Economics*, 44(3), 311-317.
- von Cramon-Taubadel, S. (1998). Estimating asymmetric price transmission with the error correction representation: an application to the German pork market, *European Review of Agricultural Economics*, 25(1), 1-18.
- Wolffram, R. (1971). Positivistic measures of aggregate supply elasticities: some new approaches - some critical notes, *American Journal of Agricultural Economics*, 53(2), 356-359.

Table 1: Tests of Non-Cointegration

	Engle-Granger	TAR ($\tau=0$)	M-TAR ($\tau=0$)	M-TAR ($\tau \neq 0$)
ρ_1	-0.208 (-2.43)	-0.214 (-1.87)	-0.416 (-2.51)	-0.606 (-3.28)
ρ_2	-	-0.203 (-1.84)	-0.164 (-1.83)	-0.165 (-1.97)
γ_1	0.408 (2.85)	0.408 (2.82)	0.473 (3.19)	0.528 (3.63)
γ_2	-0.160 (-1.12)	-0.160 (-1.11)	-0.150 (-1.07)	-0.119 (-0.87)
γ_3	0.274 (1.92)	0.274 (1.90)	0.376 (2.39)	0.348 (2.50)
τ	-	0	0	0.036
AIC	-5.66	-5.66	-5.71	-5.78
t-Max	-	-1.84	-1.83	-1.97
Φ	-	2.89	4.08	6.13
$\rho_1=\rho_2$	-	0.01 [0.93]	2.13 [0.15]	5.78 [0.02]
χ^2_{corr}	2.72 [0.60]	5.58 [0.23]	3.97 [0.41]	1.95 [0.74]
χ^2_{ARCH}	0.38 [0.98]	0.37 [0.98]	0.34 [0.99]	0.27 [0.99]
χ^2_{het}	6.46 [0.69]	10.16 [0.68]	15.29 [0.29]	11.04 [0.61]
E_p	23.12 [0.00]	23.16 [0.00]	24.06 [0.00]	24.82 [0.00]

Notes: 1. t-statistics in parentheses; and
2. p-values in square brackets.