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Market Failure and Japanese Farmland Rents

P.J. Dawson

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

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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 \(t\), \(\alpha\) 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 unbiasness when the rent-price elasticity is unity (Kellard, 2002). 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).

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 $p_2$:

$$ p_1 = \beta_0 + \beta_1 p_2 $$  

(1)

where $\beta_0$ and $\beta_1$ 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 $\beta_1$ is the same value for an increase in $p_2$ as for a corresponding fall, whereas magnitude asymmetry is where $\beta_1$ 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), Wolffram (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:

\[ r_t = \beta_0 + \beta_1 p_t + u_t \quad t=1,\ldots,n \]  

(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

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

\[
\Delta u_t = \rho_1 u_{t-1} + \sum_{i=1}^{s-1} \gamma_i \Delta u_{t-i} + v_t
\]  

(3)

where \( u_t \) are the errors in (2) and \( v_t \) is a white noise error term. Lagged values of \( \Delta u_t \) 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:

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

(4)

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

\[
I_t = \begin{cases} 
1 & \text{if } u_{t-1} \geq \tau \\ 
0 & \text{if } u_{t-1} < \tau 
\end{cases}
\]  

(5)
and \( \tau \) 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:

\[
I_t = \begin{cases} 
1 & \text{if } \Delta u_{t-1} \geq \tau \\
0 & \text{if } \Delta u_{t-1} < \tau 
\end{cases}.
\] (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 \( \tau \) 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 \( \Delta u_t \), and \( \tau \) 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 \( \tau \), 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, \( \tau \), 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 \( \rho_1 \) and \( \rho_2 \) as t-Max, the null is either that \( \rho_1 = 0 \) or that
$\rho_2=0$. Second, the $\Phi$-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 $\Phi$-statistic has more power and is preferred.

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

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

Lagged values of $\Delta r_t$ and $\Delta p_t$ are added to (7) to ensure that the error term, $\epsilon_t$, approximates white noise. In (7), $\alpha_1$ 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, $\alpha_2$ 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:

\[
\begin{align*}
    r_t &= -6.694 + 1.186p_t + \hat{u}_t \\
    R^2 &= 0.76, \quad D.W. = 0.29
\end{align*}
\]  

(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 $\rho_1$ and $\rho_2$ 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 $\rho_1$ and $\rho_2$ again imply convergence. The conclusions from the $\Phi$-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 $\tau$. A search grid yields $\tau=0.036$, and point estimates of $\rho_1$ and $\rho_2$ imply convergence. The conclusions from the $\Phi$-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, $\tau$, best fits the data. 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 $\chi^2_{corr}$ (Greene, 2012, pp.316-317), Engle's test for autoregressive conditional heteroscedasticity of order four by $\chi^2_{ARCH}$ (Greene, pp.976-977), White's test for heteroscedasticity by $\chi^2_{het}$ (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

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:

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

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

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

(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 $\Delta p_{t-1}$ do not enter the $\Delta r_t$-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 $\Delta p_t$ 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 $\chi^2_1 = 0.17$ (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 $\alpha_1$ and $\alpha_2$ in (9) and are illustrated in Figure 2. For positive discrepancies from long-run equilibrium when $\hat{\Delta}u_{t-1}\geq 0.036$, 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 $\hat{\Delta}u_{t-1}<0.036$, only 15% of the adjustment to long-run equilibrium takes place in the first year, 49% of full adjustment is complete by the end of eighth year.

\(^7\)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 $\chi^2_4 = 0.06$ (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.\textsuperscript{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 unbiasness, 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 $\Delta r_t - \Delta p_{t-1} \geq 0.036$. 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.

\textbf{Figures 2 and 3 about here}

5. \textbf{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.\textsuperscript{8} A \textit{caveat} is that adjustment coefficients have poor small-sample properties (Hansen, 1997; and Enders et al. (2007)).
and they support the rent-formation model. 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

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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


Table 1: Tests of Non-Cointegration

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<tr>
<th></th>
<th>Engle-Granger</th>
<th>TAR (τ=0)</th>
<th>M-TAR (τ=0)</th>
<th>M-TAR (τ≠0)</th>
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<td>0.408</td>
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<td>(3.19)</td>
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Notes: 1. t-statistics in parentheses; and 2. p-values in square brackets.
Figure 1: Farmland Rents and Prices in Japan (1955-2010)
Figure 2: Adjustment of Rents to Shocks in Price

- \( \Delta u(t-1) \geq 0.036 \)
- \( \Delta u(t-1) < 0.036 \)
Figure 3: Changes in the Estimated Residuals, $\Delta \hat{\mu}_t$