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Abstract

Based on newly constructed prefectural land price data, we estimate long-run equilibrium relationships using a panel cointegration analysis, and then estimate an error-correction model (ECM) for land prices. The panel cointegration analysis reveals that the PVR cum price expectation can be regarded as a long-run equilibrium relationship. The ECM finds that deviations from the long-run equilibrium and non-performing loans in particular have sizable effects on land prices. Moreover, recent regional discrepancies in land prices are closely related to deviations from the long-run equilibrium.

JEL Classification Number: G12, C23

Keywords: land price, present value relation, panel cointegration

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1 Introduction

This paper searches for the long-run equilibrium relationship for prefectural land prices, and uses the results to find determinants of land prices in Japan. The recently developed techniques of panel cointegration analysis are used.

Although cointegration analysis is a standard tool for seeking statistically the long-run equilibrium relationships among variables of interest, only a few studies have applied it to land prices in Japan, including Idee (1992), Imagawa (2002), and Yoshioka (2002). More, but still few studies are available for the U.S. property prices—see Meese and Wallace (1994), Clayton (1997), Malpezzi (1999), Meen (2002), and Gallin (2003).

Insufficient degrees of freedom seem to prevent researchers from conducting cointegration analyses of land prices. Land prices are mostly low frequency data; only annual or biannual data is available in Japan, and aggregated data typically provides a very small number of sample observations. Cointegration analysis generally requires a large number of observations, without which it is difficult for researchers to tell whether the residuals of the estimated relationships are mean-reverting in the long run.

Panel data of prefectural land prices is thought to enable us to overcome this small sample problem. Although time-series variations in data are limited, inclusion of cross-sectional variations from 47 prefectures in Japan might compensate for the insufficient degrees of freedom, within the framework of panel cointegration analysis.

Panel data of prefectural land prices could also enable us to analyze regional discrepancies. Regional discrepancies in land prices occasionally become an issue of public concern in Japan. Recently, a discrepancy has been perceived between urban and non-urban areas, in that urban land prices, especially in Tokyo, decline more slowly, whereas non-urban land prices fall more rapidly. The cross-sectional variations in our panel data might shed some light on this issue.

The paper is organized as follows. Section 2 constructs a chain-type index of prefectural land prices and shows the regional discrepancies in it. Section 3 derives theoretically equilibrium relationships from the no-arbitrage condition, and test statistically whether the data support these relationships as the long-run equilibrium. Section 4 estimates an error-correction model (ECM) of land prices to see what factors lie behind the regional discrepancies. Section 5 summarizes the main findings. Appendix A describes additional results of unit-root tests, and Appendix B details the derivation of the effective property tax rates.

2 Prefectural Land Prices

We construct panel data for the prefectural land price index from the weighted average of changes in unit prices (yen per square meter) of individual plots; the weights come from values (unit prices multiplied by their sizes) in previous years. Since we revise the weights every time period, the resulting series comprises a chain-type index.

Specifically, the land price P_{it} in prefecture i in period t is calculated as:

$$\Delta p_{it} = \sum_{j \in i} \frac{V_{j,t-1}}{\sum_{j \in i} V_{j,t-1}} \Delta p_{jt},$$

where P_{jt} is the unit price of plot j in prefecture i and V_{jt} is its value. Here p_{it} and p_{jt} are the natural logarithms of P_{it} and P_{jt}. Δ denotes the first difference operator. P_{jt} is obtained from the Officially Published Land Prices (OPLP) of the Ministry of Land, Infrastructure and Transport (MLIT).[1]

We call the resulting series the Weighted-average Officially Published Land Price index (W-OPLP), to distinguish it from the OPLP. In fact, as part of the OPLP, the MLIT compiles prefectural land prices as the *simple* average of changes in unit prices. We believe, as has been widely pointed out, that there is no rationale for taking the simple average; the weighted average is superior to the simple average, at least for studying changes in value of land assets over a prefecture.

Unit prices of the OPLP (and hence the W-OPLP) are appraisal-based prices, which can differ substantially from transaction-based prices; appraisal-based prices tend to be less volatile and to lag behind transaction-based prices (Nishimura and Shimizu, 2002; Saita, 2003). Although we are aware of this problem, as far as prefectural land prices are concerned there is no alternative to using appraisal-based prices since the other prefectural land prices based on SNA statistics also utilize appraisal-based prices. However, if the appraisal-based prices track the transaction-based prices in the long run, then panel cointegration analysis is relatively free from this problem, since it is the long-run relationship that matters for cointegration.

Some researchers (Idee, 1997; Kousai, Itou, and Sadamoto, 1999; Fujiwara and Shinke, 2003) have used the SNA based prefectural land prices, but we prefer the W-OPLP for two reasons. First, we observe some peculiarities in the SNA based prefectural land prices: land prices in a considerable number of prefectures (Aomori, Yamagata, Fukui, Shimane, Tokushima, Kochi, Saga etc.) rose sharply in the latter half of the 1990s.[2] We believe this is scarcely supported by other statistics or anecdotal evidence. Second, we believe the long delay in publication unsuitable for policy analysis—as of June 2004, the SNA based prefectural land prices are available only up to 2001.

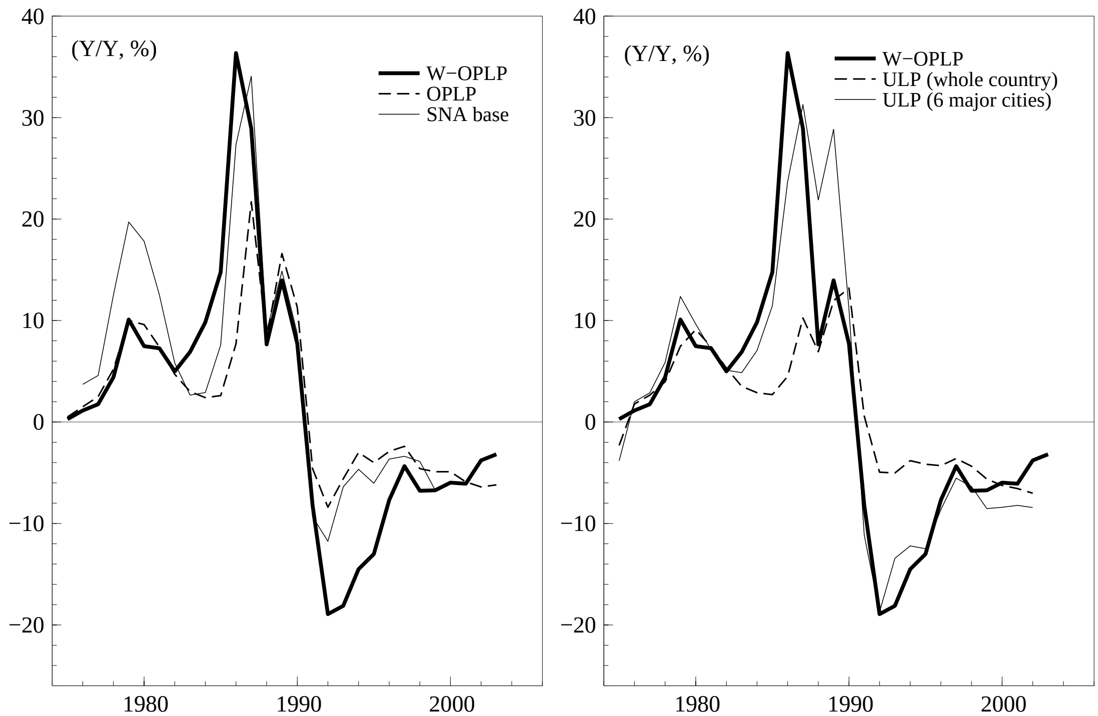
In aggregate, for whole of the country, changes in the W-OPLP resemble those in the Urban Land Price (ULP) index[3] (six major cities) and the SNA based land price (Figure

[1]At January 1, 2004, there were 31,866 surveyed plots, of which up to 3,254 plots were in Tokyo. Newly surveyed plots were added to the sample each year partly in exchange of plots dropped from the sample. These new plots are omitted from the calculation of P_{it} (but included for $P_{i,t+1}, \ldots$), as $P_{j,t-1}$ is not available.

[2]The SNA based prefectural land prices are derived as the values of land assets of prefectures (Cabinet Office, "Annual Report on National Account") divided by their areas excluding farmland and forests (Ministry of Public Management, Home Affairs, Post and Telecommunications (MPHPT), "Summary Report on Prices of Fixed Assets").

[3]Urban Land Price index is another appraisal-based land price index compiled by the Japan Real

Figure 1: Comparison of Land Price Measures



Note: Although the OPLP and the W-OPLP are appraised at the beginning of the year, in the figure and hereafter they are treated as data at the end of the previous year.

1). This is because the unit prices of the properties are higher in the six major cities, where both the W-OPLP and the SNA based land prices put heavier weights.

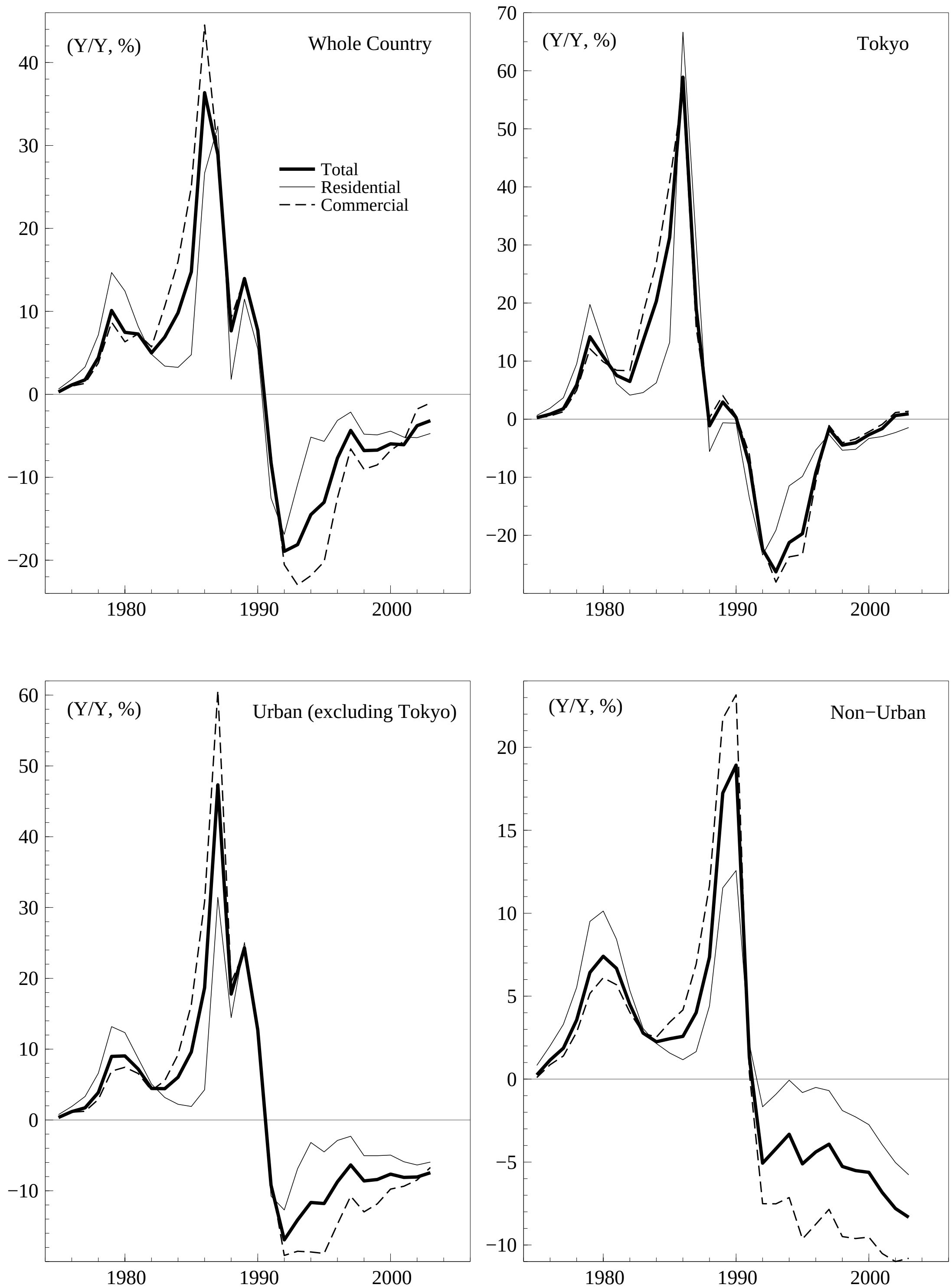
Figure 2 shows the W-OPLP and its breakdowns by region (Tokyo/Urban excluding Tokyo/Non-Urban) and by usage (residential/commercial). We define urban prefectures as those in which government-appointed 12 major cities (*seirei shitei toshi*) are located: Hokkaido, Miyagi, Saitama, Chiba, Tokyo, Kanagawa, Aichi, Kyoto, Osaka, Hyogo, Hiroshima and Fukuoka.

Regional discrepancies are clearly seen in the figure, as follows:

- In Tokyo (upper right panel), land prices rocketed in 1985-1986. After the bubble burst in the early 1990s, prices plummeted by more than 25 percent on an year on year basis. In the latter half of the 1990s, the land price declined more slowly and began to increase from 2002. In 2003, it rose by nearly one percent.

Estate Institute. Although it has a breakdown of six major cities (Tokyo, Yokohama, Nagoya, Kyoto, Osaka and Kobe), no prefectural breakdown is available, nor unit prices of individual plots.

Figure 2: W-OPLP



- In the other urban prefectures (lower left panel), the land price shot up in 1987, just after the rally in Tokyo. In the early 1990s it declined by more than 15 percent. In the latter half of the 1990s it declined rather more slowly, but still fell by more than five percent even in 2003.
- In the non-urban prefectures (lower right panel), land price did not increase sharply until 1989. Both the hike and the fall during and after the bubble period were relatively small in magnitude. Contrary to the urban prefectures, there is no indication of slowing down in its declining speed. In 2003, it dropped by nearly eight percent.

In the recent period, Figure 3 confirms the regional discrepancies in a scatter diagram. The vertical axis shows acceleration/deceleration of changes in the W-OPLP from 1995, and the horizontal axis shows gross prefectural expenditures in 1995. Prefectures with larger expenditures, which tend to be urban prefectures such as Tokyo, experience the more deceleration of fall in land prices. Prefectures with smaller expenditures, which tend to be rural prefectures, suffer from an accelerating fall in land prices.

Below, we investigate the factors behind these regional discrepancies by estimating an ECM for land prices, which explicitly takes into account the long-run equilibrium relationship.

3 Long-Run Equilibrium

3.1 PVR

Following Meese and Wallace (1994) and Clayton (1997), we test whether the present value relation (PVR) for land prices holds in the long run.

Assuming risk-neutrality, the PVR can be derived from:

$$Y_{it} + (P^e_{i,t+1} - P_{it}) = r_{it} P_{it}, \tag{1}$$

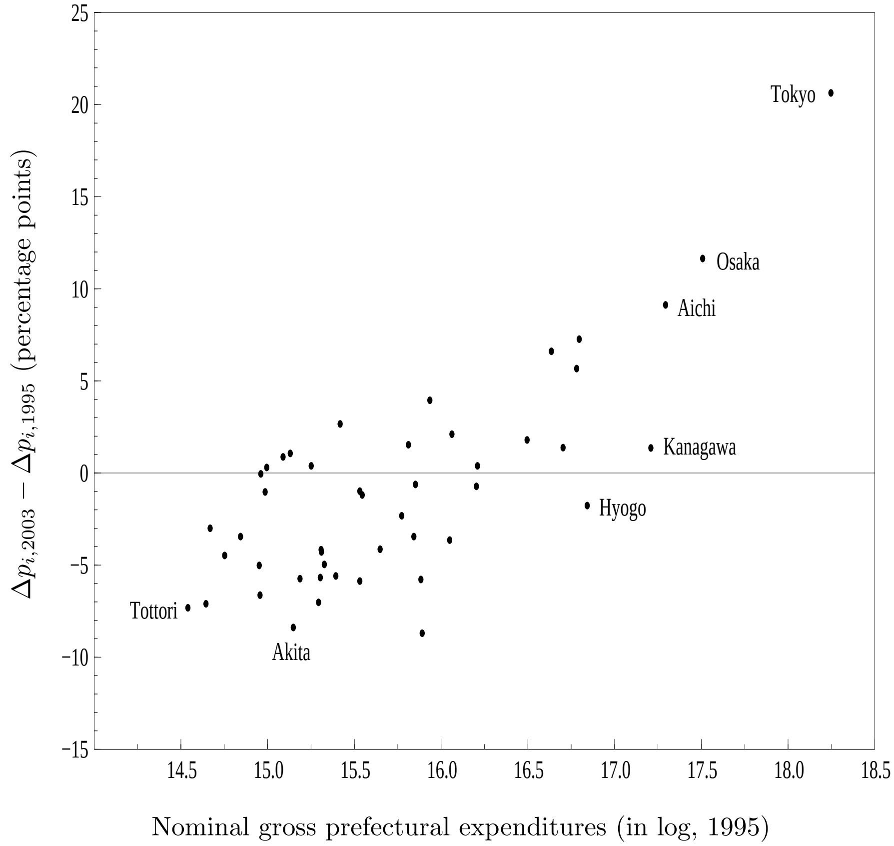
where Y_{it} is the nominal rental income, $P^e_{i,t+1}$ is the expected price in period $t+1$, r_{it} is the cost of capital, which consists of the nominal interest rate i_t and the property tax rate τ_{it} ($r_{it} = i_t + \tau_{it}$).

Equation (1) is the no-arbitrage condition for the land market. The left-hand-side of the equation expresses the fact that the expected return for property investment is the sum of income gain (Y_{it}) and capital gain ($P^e_{i,t+1} - P_{it}$). The right-hand-side indicates that the cost of capital for the investment comprises interest and tax payments. If the no-arbitrage condition holds, there should be no discrepancy between the expected return and its cost, as implied in equation (1).

It is well known in the asset pricing model that this relation does not preclude a bubble in land prices. Equation (1) can be rewritten as:

$$P_{it} = \frac{Y_{it} + P^e_{i,t+1}}{1 + r_{it}}.$$

Figure 3: Land Prices and Gross Prefectural Expenditures



Recursive substitution of $P^e_{i,t+1}$, $P^e_{i,t+2}$,... yields

$$P_{it} = E_t \left[\sum_{h=1}^{\infty} \left\{ \prod_{k=1}^{h} \left(\frac{1}{1 + r_{i,t+k}} \right) \right\} Y_{i,t+h} + \lim_{h \to \infty} \prod_{k=1}^{h} \left(\frac{1}{1 + r_{i,t+k}} \right) P_{i,t+h} \right], \tag{2}$$

where E_t is the expectation operator conditional on the information set Ω_t ($E[.|\Omega_t]$) and $P^e_{i,t+h} = E_t[P_{i,t+h}]$ for $h = 1, ...$

To preclude the bubble, it is necessary to assume that the land price P_{it} does not grow faster than the capital cost r_{it} (transversality condition), so that the last term of equation (2) vanishes:

$$E_t \left[\lim_{h \to \infty} \prod_{k=1}^{h} \left(\frac{1}{1 + r_{i,t+k}} \right) P_{i,t+h} \right] = 0.$$

In addition, if we assume static expectation, so that the future capital cost $r_{i,t+k}$ is same as its current value r_{it}, and rental income Y_{it} grows at the constant rate g^e_{it}, equation

(2) gives the well-known fundamental price relation:

$$P_{it} = \frac{Y_{it}}{r_{it} - g_{it}^e}. \tag{3}$$

Below, we use panel cointegration tests to statistically examine whether the PVR in the form (1) or (3) holds in the long run. As surveyed by Inoue, Idee, and Nakagami (2002), the short-run PVR is often empirically rejected. This is probably because the land market does not conform to the usual assumptions of efficient market theory, which entails no chance of arbitrage. Instead, the land market is characterized by information asymmetries, high transaction costs and heterogeneous assets and beliefs. The violation of these assumptions might be unimportant in the longer term, however.

Below, we shall estimate an empirical correspondent to equation (1):

$$p_{it} = \alpha p_{i,t+1}^e + \beta y_{it} - \gamma r_{it} + d_t + \eta_i + \nu_{it}, \tag{4}$$

where p_{it}, p_{it}^e and y_{it} denote the logarithms of P_{it}, P_{it}^e and Y_{it}. The first three terms correspond to a linear approximation of equation (1) (Campbell and Shiller, 1988a,b):

$$p_{it} \simeq \rho p_{i,t+1}^e + (1-\rho) y_{it} - r_{it} + \kappa, \tag{5}$$

where ρ and κ are constants.[4] We do not impose the theoretical restrictions that $\alpha + \beta = 1$ and $\gamma = 1$ from the outset. The remaining terms are disturbance terms, which are supposed to consist of time specific effects d_t, individual effects η_i, and idiosyncratic shocks ν_{it}.

We likewise estimate an empirical relation corresponding to the fundamental price relation (3) as:

$$p_{it} = \phi y_{it} - \psi \ln(r_{it} - g_{it}^e) + d_t + \eta_i + \nu_{it}, \tag{6}$$

without imposing theoretical restrictions of $\phi = \psi = 1$.

3.2 Panel Cointegration Test

In this section we conduct panel cointegration tests to see whether the two variants of the PVR above can be regarded as the long-run equilibrium, by checking their stationarity.

[4]Without the prefectural subscript i, equation (1) can be written as:

$$\frac{Y_t + P_{t+1}^e}{P_t} = 1 + r_t.$$

On taking the logarithm, the RHS is $\ln(1 + r_t) \simeq r_t$. Using $\delta_t = \ln(Y_{t-1}/P_t) = y_{t-1} - p_t$, the LHS becomes

$$h_t = \ln(\exp(\delta_t - \delta_{t+1}) + \exp(\delta_t)) + \Delta y_t.$$

The above linear approximation can be derived from the first order Taylor expansion of $h_t(\delta_t, \delta_{t+1})$ around $h_t(\delta, \delta)$, where δ is the long-run equilibrium of δ_t, $\rho = 1/(1 + \exp(\delta))$, and $\kappa = \ln(1 + \exp(\delta)) - \delta \exp(\delta)/(1 + \exp(\delta))$.

Table 1: Panel Unit Root Test (Hadri)

	Level	First Difference
p	11.21** (0.00)	0.97 (0.17)
y	21.06** (0.00)	1.18 (0.12)
r	9.63** (0.00)	–2.03* (0.02)

Note: "**" and "*" denote statistical significance at the 1% and 5% levels, respectively. Figures in parentheses are the p-value. Two lags are used for kernel estimation.

The data used are as follows: (i) The prefectural land price P_{it} is the above W-OPLP. (ii) The nominal rental income Y_{it} is approximated by gross prefectural expenditures (Cabinet Office, "Annual Report on Prefectural Accounts").[5] (iii) The nominal interest rate i_t is obtained from the "Average Contracted Interest Rates on Loans and Discounts," (Bank of Japan). (iv) The property tax rate τ_{it} is calculated from the effective tax rates of Real Property Acquisition Tax, Registration and License Tax, Property Tax, Urban Planning Tax, and Land Value Tax (see Appendix B for details). (v) The expected growth rate g_{it}^e is the moving backward average of nominal growth ($\sum_{k=1}^{3} \Delta y_{i,t-k}/3$). (vi) The expected land price $P_{i,t+1}^e$ is either the actual price (perfect foresight) or a prediction made by a univariate time-series model—after several experiments, we choose ARIMA(2,1,0) estimated by pooled regression, but the following results are robust against alternative specifications.

Prior to panel cointegration tests, we check the degree of integration of each variable using the panel unit-root tests of Hadri (2000). To remove cross-sectional correlation, we subtract the average of each variable, as $x_{it} - (1/N)\sum_{i=1}^{N} x_{it}$ where x_{it} is a variable in question.

Table 1 summarizes the results. For all three variables in level, the null hypothesis that "x_{it} do not have unit roots for all prefectures" is rejected. In first difference, the null hypothesis is not rejected for p_{it} and y_{it}, so that these variables are I(1), meaning integrated of order one. For r_{it} the null hypothesis is rejected at the five percent significance level, so that the variable is possibly more than I(2). However, since both the ADF test and the Fisher test in Appendix A strongly suggest that r_{it} is I(1), we assume it is I(1) and proceed to the panel cointegration test with the caveat.

We estimate cointegrating vectors in equations (4) and (6) by the Group-Mean Fully Modified OLS (FMOLS) of Pedroni (2000, 2001). There are three steps: (A) we first eliminate individual effects η_i using a within-group transformation, such as $p_{it}^{WG} = p_{it} - (1/T)\sum_{t=1}^{T} p_{it}$; (B) next, for each prefecture, we estimate parameters by FMOLS (Phillips and Hansen, 1990) from these within-group transformed variables; (C) we finally obtain

[5] Indeed, this is a rough proxy of rental income, but an alternative measure such as prefectural outputs of real estate industry plus imputed housing rents yields similar results.

cointegrating vectors by taking the average of these parameters across individual prefectures (group-mean).

Group-mean regression gives a different interpretation to the alternative hypothesis of the t-test from pooled regression (i.e., instead of steps (B) and (C) above, the relevant parameters are estimated from pooled data). For example, consider the t-test on α in equation (4), which has the null hypothesis H_0: $\alpha_i = 0$ for all i. In the case of group-mean regression, the alternative hypothesis becomes H_1: $\alpha_i \neq 0$. But in the case of pooled regression, the alternative hypothesis becomes H_1': $\alpha_i = \alpha_A \neq 0$. The group-mean regression has a less restrictive alternative hypothesis, since it does not require that α_i take a common value α_A (see also footnote 8 below).

Since Group-Mean FMOLS assumes no cross-sectional correlation across individual residuals, we control common shocks by including time dummies in the Group-Mean FMOLS regressions. This is equivalent to subtraction of cross-sectional means ($x_{it} - (1/N)\sum_{i=1}^{N} x_{it}$) in the above Hadri test.

Based on samples from 47 prefectures from 1976 to 2001, we obtain the following results of estimation (two lags are used for kernel estimation):

- PF: Equation (4) with perfect foresight,

$$\begin{array}{llll} p_{it}^* & = 0.86 p_{i,t+1}^e & + 0.20 y_{it} & - 3.27 r_{it}, \\ & \quad (50.7) & \quad (3.04) & \quad (5.12) \end{array} \tag{7}$$

panel ρ: –0.72, panel PP: –4.07, panel ADF: –4.86, panel ν: 4.90,
group ρ: 0.79, group PP: –4.53, group ADF: –6.44.

- ARIMA: Equation (4) with ARIMA prediction,

$$\begin{array}{llll} p_{it}^* & = 0.88 p_{i,t+1}^e & + 0.15 y_{it} & - 0.58 r_{it}, \\ & \quad (88.9) & \quad (3.53) & \quad (1.35) \end{array} \tag{8}$$

panel ρ: –0.72, panel PP: –3.27, panel ADF: –7.48, panel ν: 3.95,
group ρ: 1.80, group PP: –2.13, group ADF: –8.61.

- FM: Fundamental price of equation (6),

$$\begin{array}{lll} p_{it}^* & = 0.91 y_{it} & - 1.67 \ln(r_{it} - g_{it}^e), \\ & \quad (9.85) & \quad (17.2) \end{array} \tag{9}$$

panel ρ: 3.77, panel PP: 4.67, panel ADF: 3.58, panel ν: –2.18,
group ρ: 5.83, group PP: 6.76, group ADF: 5.05.

Figures in parentheses are t-values. Seven statistics (panel/group ρ, panel/group PP, panel/group ADF, and panel ν) are cointegration tests of Pedroni (1999), with the null hypothesis of no cointegration.[6] One side of the standard normal distribution is used for rejection region; the upper-side is used for the variance-ratio test (panel ν), and the lower-side is used for the other tests. As in Phillips-Perron tests, panel/group ρ, panel/group PP, and panel ν make use of non-parametric corrections for autocorrelated error terms, and panel/group ADF correct parametrically; see Pedroni (1999) for more details.

In the case of perfect foresight and in the case of ARIMA prediction, the PVR cum price expectation, which corresponds to equation (1), yields generally reasonable results. Estimated coefficients in both equations (7) and (8) are significant in most cases, and signed as theory predicts. Parameter values, especially those of equation (8), are not far from theoretical prediction: the restriction $\alpha + \beta = 1$ is not rejected in either case (t-statistics on the restriction are 0.47 and 0.02 for equations (7) and (8), respectively), and the restriction $\gamma = -1$ is not rejected for the latter case (t-statistics on the restriction are 3.35 and 1.42, respectively). Finally, except for panel ρ tests (and the group ρ test in the case of equation (7)), all the cointegration tests reject the null hypothesis of no cointegration.

The fundamental price relation, which corresponds to equation (3), nevertheless does not appear promising. Although the estimated coefficients in equation (9) are statistically significant and correctly signed, all the cointegration tests fail to reject the null hypothesis of no cointegration.

To check the robustness of the estimated cointegrating vectors, we further estimate the same equations by the Group-Mean Dynamic OLS (DOLS) method of Pedroni (2001). We obtain parameters for individual prefectures by the following dynamic regressions in place of FMOLS (in step (B) above):

$$p_{it} = \alpha' p^e_{i,t+1} + \beta' y_{it} - \gamma' r_{it} + \sum_{j=1}^{2} \varrho_j \Delta p^e_{i,t-j} + \sum_{j=-1}^{1} \varpi_j \Delta y_{i,t-j} + \sum_{j=-1}^{1} \varsigma_j \Delta r_{i,t-j}, \text{ for (4).}$$

$$p_{it} = \phi' y_{it} - \psi' \ln(r_{it} - g^e_{it}) + \sum_{j=-1}^{1} \varpi_j \Delta y_{i,t-j} + \sum_{j=-1}^{1} \varsigma_j \Delta \ln(r_{i,t-j} - g^e_{i,t-j}), \text{ for (6).}$$

The estimations yield similar cointegrating vectors, backing up the results obtained by FMOLS.

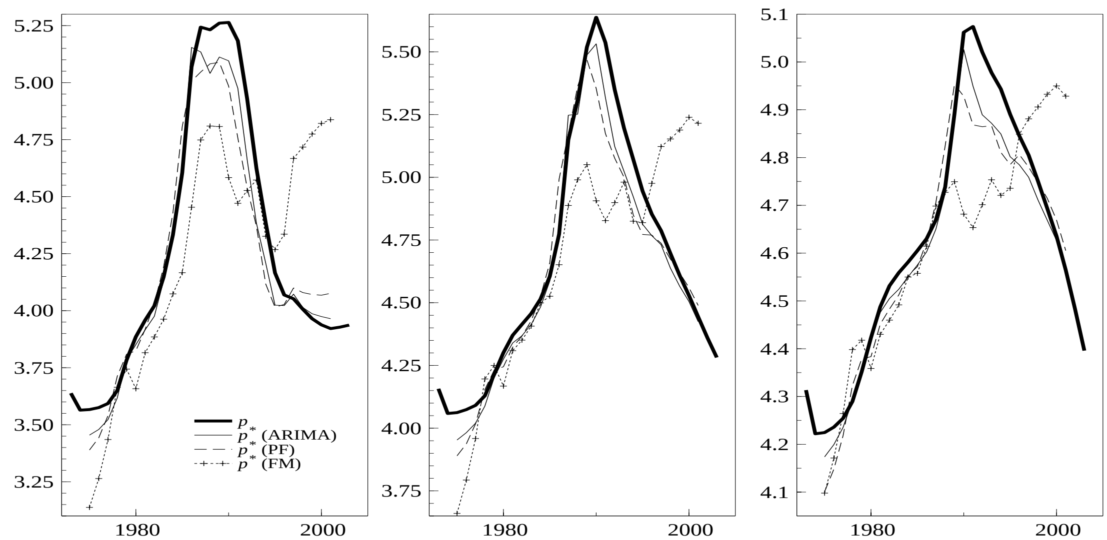
- PF: Equation (4) with perfect foresight,

$$\begin{array}{rlll} p^*_{it} & = 0.79 p^e_{i,t+1} & +0.53 y_{it} & -4.47 r_{it}. \\ & \quad (61.6) & (8.02) & (7.79) \end{array}$$

[6]Suppose we regress the estimated error $\hat{\nu}_{it}$ in equations (4) and (6) on its own lag $\hat{\nu}_{i,t-1}$ and obtain the coefficient ρ^ν_i. The null hypothesis of all the tests is "no cointegration relationship for all prefectures (H_0: $\rho^\nu_i = 1$ for all i)." The alternative for panel ρ/PP/ADF/ν is "there is a cointegration relationship and the degree of residual autocorrelation is same across prefectures (H_1: $\rho^\nu_i = \rho^\nu < 1$ for all i)," whereas the alternative for group ρ/PP/ADF is "there is a cointegration relationship and the degree of residual autocorrelation may differ across prefectures (H'_1: $\rho^\nu_i < 1$ for all i)."

Figure 4: Cointegration Relationships

(1) Tokyo (2) Urban other than Tokyo (3) Non-Urban



Note: p^* (PF), p^* (ARIMA) and p^* (FM) are calculated from equations (7), (8) and (9), respectively. Match means and ranges.

- ARIMA: Equation (4) with ARIMA prediction,

$$\begin{array}{rcccc} p^*_{it} & = & 0.82p^e_{i,t+1} & +0.32y_{it} & -2.58r_{it}. \\ & & (118) & (8.82) & (7.68) \end{array}$$

- FM: Fundamental price of equation (6),

$$\begin{array}{rccc} p^*_{it} & = & 1.03y_{it} & -1.82\ln(r_{it} - g^e_{it}). \\ & & (21.7) & (30.1) \end{array}$$

The results of the above cointegration tests are made clear in Figure 4, in which p^* calculated from equations (7)-(9) is plotted together with actual land price p. Both p^* (PF) and p^* (ARIMA), which are two variants of the PVR cum price expectation, track the actual price development. In contrast, p^* (FM), the fundamental price, diverges from p starting from the latter half of the 1990s. Since cointegration requires that p^* follow p in the long run, it is clear that both p^* (PF) and p^* (ARIMA) are cointegrated with p, whereas p^* (FM) is not.

The fact that the fundamental price relation is not supported by data as a long-run equilibrium indicates that price expectation, which allows for the possibility of a bubble in land prices, has played a vital role during this sample period. The failure to find a cointegration relationship (equation (9)) could be due to inadequacy of the assumption of static expectation and/or risk neutrality. However, since the PVR cum price expectation under the same risk neutrality assumption is supported by data as a long-run equilibrium, it appears that price expectation is indispensable. About half the sample period 1976-2001 is characterized by episodes of emergence of the bubble (the latter half of the 1980s) and its bursting (the 1990s), so that if the bubble endured during a significant part of the sample period, price expectation cannot be neglected.

4 Error Correction Model

In this section, we estimate an ECM of land prices by embedding the long-run equilibrium relationship of equation (8). Equation (7) is not used, because we believe that the assumption of perfect foresight is less realistic.

The estimated ECM specification is

$$\Delta p_{it} = -\theta(p - p^*)_{i,t-1} + \lambda \Delta z_{it} + \varepsilon_{it}, \tag{10}$$

where Δz_{it} are explanatory variables which determine the short-run dynamics, and ε_{it} is the disturbance term.

For Δz_{it} we consider the following six variables, whose developments are shown in Figure 5 by region.

1. Δr_{it}: First difference of the capital cost r_{it}.

   In Figure 5, r_{it} in Tokyo was very low in the latter half of the 1990s. This reflects the low effective property tax rates. At that time, valuation of the tax base did not fully keep up with the hike in actual land prices.

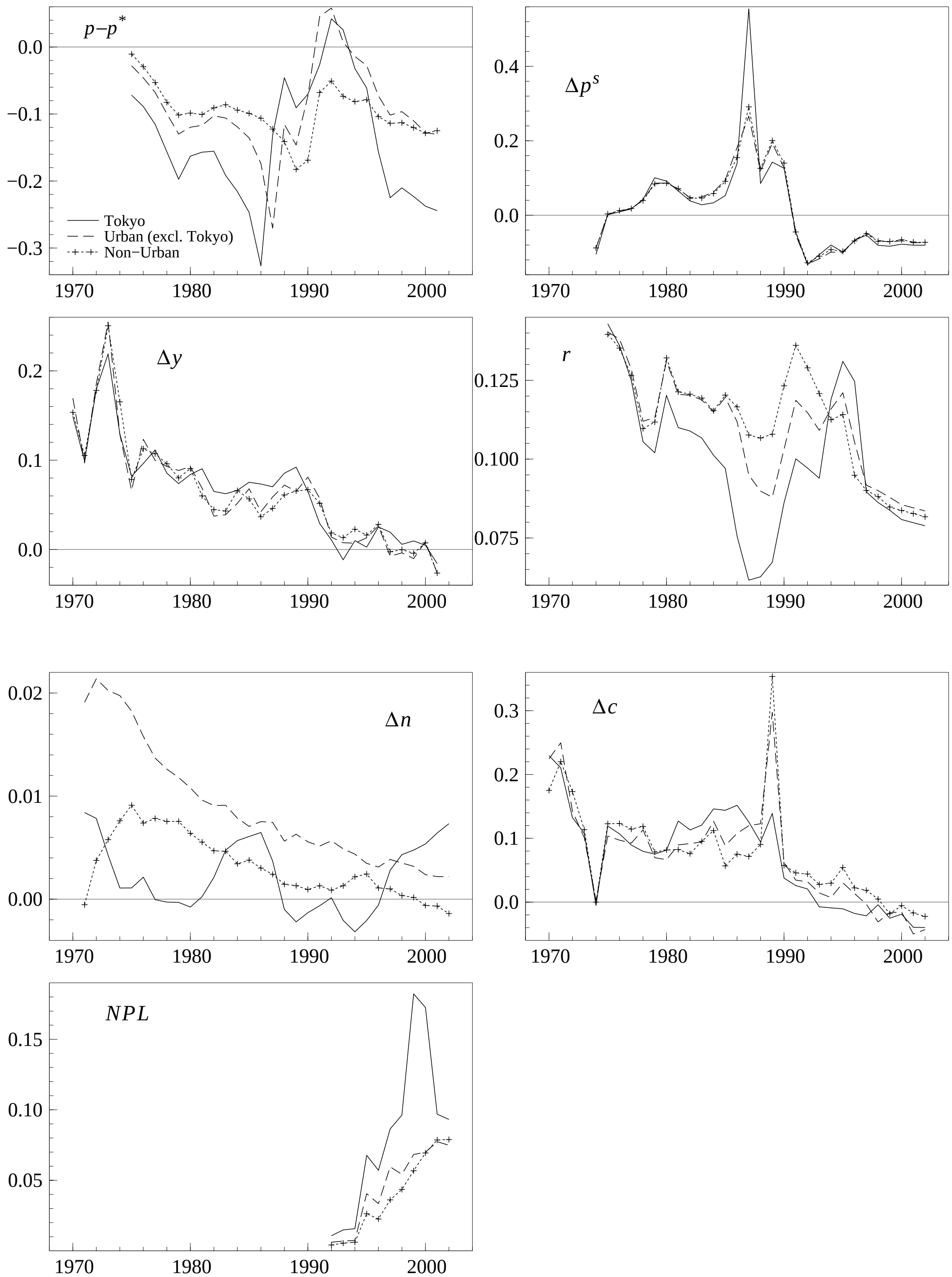
   Despite the easing monetary policy and the tax relief on the Real Property Tax and the Registration and License Tax, r_{it} rose in Tokyo in the first half of the 1990s. This is because the effective rates of the Property tax increased, since the evaluation price for tax base was not fully adjusted to the collapse of land prices. Also, r_{it} declined in non-urban areas, where the fall in land price was smaller at that time, and the discrepancy between tax-base evaluation prices and actual land prices was not so large. In 1997, r_{it} fell even in Tokyo, once adjustment was made for the evaluation price.

2. Δy_{it}: First difference of the nominal prefectural income y_{it}.

3. Δn_{it}: Change in the number of residents in each prefecture (MPHPT, "Basic Residential Registers"). The variable is included to account for demographic effects.

   The number of residents in Tokyo shows a cyclical fluctuation (Figure 5). It increased around the middle of the 1980s, and declined from the late 1980s to the middle of the 1990s; recently it increased again, a phenomenon known as the "Return to Tokyo." In contrast, the population growth in the non-urban areas shows a declining trend, and turns out to be negative from 2000.

Figure 5: Explanatory Variables



4. $\Delta p^s_{i,t-1}$: Change in land prices in neighboring prefectures in the previous year. The variable is included to account for spatial correlation.

   Spatial correlation is often said to be important. For instance, it is claimed that, during the bubble period, the hike in land prices in Tokyo infected first other urban areas and then non-urban areas. This lead-lag relationship is evident in Figure 2.

   Δp^s_{it} is calculated as the weighted average of changes in land prices in other prefectures, where the weights are derived from distances measured by economic linkages with prefecture i. The quantity of cargos transported between relevant prefectures (MLIT "National Survey on Cargo Transportations") is used as a proxy to show how close the economic linkages are. That is,

$$\Delta p^s_{it} = \sum_h w^i_{ht} \Delta p_{ht},$$

   where w^i_{ht} is the share of prefecture h calculated from the quantity of cargo transported between prefectures i and h ($w^i_{ht} = T_{iht} / \sum_h T_{iht}$).

5. $\Delta c_{i,t-1}$: Changes in outstanding bank loans in the previous year (Bank of Japan, "Loans and Discounts Outstanding by Prefecture"). The variable is included to account for the effect of bank credits.

   Bank loans are often said to be closely related to the property bubble; banks expanded their credits, backed by appreciated land collateral, raising the land prices, increasing bank loans, and so on. This circle appears to have reversed after the bursting of the bubble: banks were forced to lend less, since the value of collateral decreased following a fall in land prices.

   In 1989, spikes are seen in Δc_{it} for urban areas other than Tokyo and for non-urban areas (Figure 5). This is mainly due to a change in the status of banks. As part of financial liberalization, *sogo* banks, which are small banks established as joint stock companies, changed their status to regional banks (Regional Bank II), whose loans then began to be included in the statistics. Since the shares of these banks are larger in prefectures other than Tokyo, the effect of the change is more obvious in these prefectures.

6. $NPL_{i,t-1}$: The non-performing loan (NPL) ratio in the previous year. The NPL ratio is calculated as risk management loans divided by total loans outstanding (Japanese Bankers Association, "Financial Statements of All Banks").

   The NPL ratio aims to capture two effects. First, it may serve as a proxy of forbearance lending. Sekine, Kobayashi, and Saita (2003) showed that, in the 1990s, banks did not write off the NPLs aggressively enough by their forbearance policy. In this case, c_{it} reflects both active and inactive loans, so that the inactive part due to forbearance lending should be adjusted. Second, it may reflect risk premium in the corresponding prefectures: banks might require explicitly or implicitly some risk premium in accordance with their NPL ratios.

The NPL ratio for each prefecture is calculated from the ratios for regional banks whose headquarters are located in the corresponding prefecture. They are subject to the following caveats:

- These NPL ratios may not fully reflect the underlying trend in Tokyo and Osaka, where regional banks have relatively smaller shares.
- Regional banks may have piled up the non-performing loans, not in the prefectures of their headquarters, but in other prefectures such as Tokyo or Osaka.
- NPL ratios are not available before 1991, since most banks began to publish the data from 1992.

Despite these caveats, it seems reasonable to try NPL_{it}. For instance, in most of the 47 prefectures in Japan, regional banks have dominant shares, so that neglect of the NPLs of city banks may not be important. Banks might require risk premium so long as their balance-sheets are hampered irrespective of where the NPLs are accumulated. According to Sekine, Kobayashi, and Saita (2003), forbearance lending became notable from the 1990s, so that lack of data prior to 1992 might not be harmful.

NPL_{it} is relatively high in Tokyo, where the property bubble was largest (Figure 5). In non-urban areas, NPL_{it} was relatively low in the 1990s, but it has increased recently to the same level as in urban areas other than Tokyo.

We estimate equation (10) as the Random Coefficients Model (Swamy, 1970) to allow for heterogeneity in parameters.[7] That is,

$$\Delta p_{it} = -\theta_i (p - p^*)_{i,t-1} + \lambda_i \Delta z_{it} + \varepsilon_{it},$$

where individual parameters follow stochastic processes as:

$$\theta_i = \theta + \xi_i, \;\; \lambda_i = \lambda + \zeta_i.$$

These parameters can deviate stochastically from common parameters θ and λ by disturbance terms ξ_i and ζ_i respectively.

Under some additional assumptions, it can be shown that the best linear unbiased estimators of θ and λ are obtained as a matrix-weighted average of the least-squares estimator for each prefecture, with weights inversely proportional to their covariance matrices.

The result of estimating equation (10) is given in column (1) of Table 2. All coefficients have the expected signs, and are statistically significant at the conventional levels, except that the coefficient on Δr_{it} is marginally insignificant. A test for parameter homogeneity (H_β) is strongly rejected, supporting a Random Coefficients Model in which the parameters are assumed to be heterogeneous.

[7] Pesaran and Smith (1995) stress the importance of taking into account parameter heterogeneity even in estimation of long-run equilibrium relationships. In fact, the Group-Mean FMOLS and DOLS in the previous section, which take the average of individually estimated coefficients, also allow for parameter heterogeneity.

Table 2: ECM Specification of Land Prices (Δp_{it})

	(1)	(2)
$(p-p^*)_{i,t-1}$	–0.64 (0.10)***	–0.65 (0.10)***
Δr_{it}	–0.41 (0.27)	–0.41 (0.26)
Δy_{it}	0.34 (0.09)***	
$\Delta \bar{y}_t$		0.42 (0.11)***
Δn_{it}	2.87 (1.11)***	2.04 (1.08)*
$\Delta p^s_{i,t-1}$	0.18 (0.06)***	0.15 (0.05)***
$\Delta c_{i,t-1}$	0.22 (0.06)***	0.22 (0.05)***
$NPL_{i,t-1}$	–0.65 (0.20)***	–0.65 (0.18)***
constant term	–0.08 (0.01)***	–0.09 (0.01)***
Sample Periods	FY1977-FY2001	FY1977-FY2002
Standard Error	0.058	0.056
Prefectures	47	47
Observations	1,175	1,222
H_β	1154.7 [0.00]	1125.3 [0.00]

Notes:

1. Estimated as a Random Coefficients Model (Swamy, 1970).
2. Figures in parentheses are standard errors. "***", "**" and "*" denote statistical significance at the 1%, 5% and 10% levels, respectively.
3. H_β is a test for parameter homogeneity (the null hypothesis corresponds to homogeneous parameters across prefectures), which follows the χ^2 distribution with $K(n-1)$ degrees of freedom (K is the number of explanatory variables, and n is the number of prefectures). Figures in squared brackets are p-values.

In column (2) of the same table, nominal GDP growth $\Delta\bar{y}_t$ is used in place of Δy_{it}, so that the sample period runs up to 2002. In Figure 5, there is no major difference in Δy_{it} across regions, which tends to support the use of $\Delta\bar{y}_t$. Estimates are almost identical to those in column (1).

From the parameters in column (2) of Table 2, we calculate the contributions of the explanatory variables to annual changes in land prices (Figure 6). We obtain an almost identical figure to Figure 6, even using the parameters in column (1). In the figure, the contributions of $(p - p^*)_{i,t-1}$ are adjusted by the estimated constant term, which corresponds to κ in equation (5).

- In Tokyo, $(p - p^*)_{i,t-1}$ exerts downward pressure in the first half of the 1990s, but upward pressure from the late 1990s. With almost the same timing, a positive contribution is observed from Δn_{it}. Although both $NPL_{i,t-1}$ and $\Delta p^s_{i,t-1}$ continue to exert downward pressure, the negative contribution of $NPL_{i,t-1}$ is smaller in 2002 than in 2000-2001, which partly explains the bottoming out of land price in Tokyo.

- In the other urban areas, the negative contribution of $(p - p^*)_{i,t-1}$ gradually diminishes, but does not become positive even in 2002. The negative contributions of $NPL_{i,t-1}$ and $\Delta c_{i,t-1}$ increase, canceling out the above diminishing negative contribution of $(p - p^*)_{i,t-1}$; as a result, land prices in this area still decline significantly. Meanwhile, Δn_{it} continues to exert upward pressure, but $\Delta p^s_{i,t-1}$ exerts downward pressure on land prices.

- In the non-urban areas, the story is similar to urban areas other than Tokyo: the diminishing negative contribution of $(p-p^*)_{i,t-1}$ is canceled out by the larger negative contributions of $NPL_{i,t-1}$ and $\Delta c_{i,t-1}$. However, because the negative contribution of $(p - p^*)_{i,t-1}$ diminishes more slowly than in non-Tokyo urban areas, the decline in land price in the non-urban area accelerates in recent years, in contrast to the almost constant declining speed in the non-Tokyo urban areas. Meanwhile, Δn_{it} began to exert slight downward pressure from 2000.

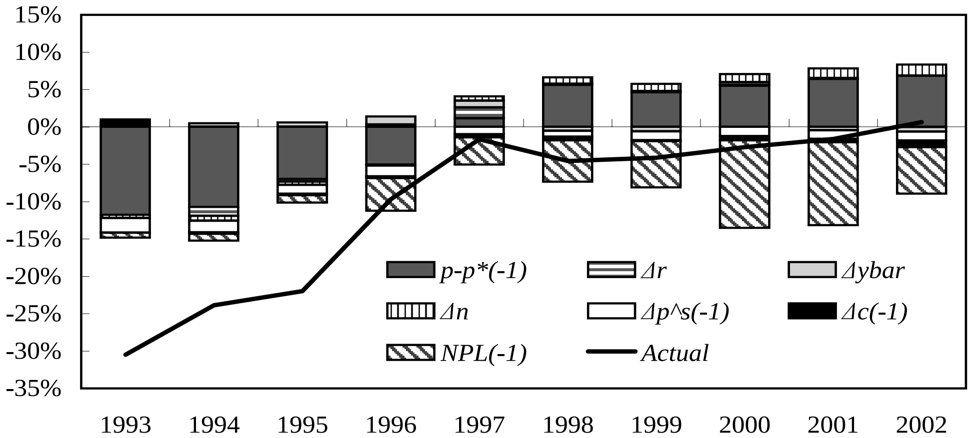
Figure 7 also confirms that the regional discrepancies shown in Figure 3 reflect the error correction mechanism regarding $(p - p^*)_{i,t-1}$: a deviation from the long-run equilibrium price. Positive correlation is evident in Figure 7 between acceleration/deceleration of land prices from 1995 to 2002 and changes in $(p - p^*)_{i,t-1}$. This implies that in urban areas, where actual land prices dropped sharply in the first half of the 1990s, those land prices become more in line with the long-run equilibrium prices.

5 Conclusion

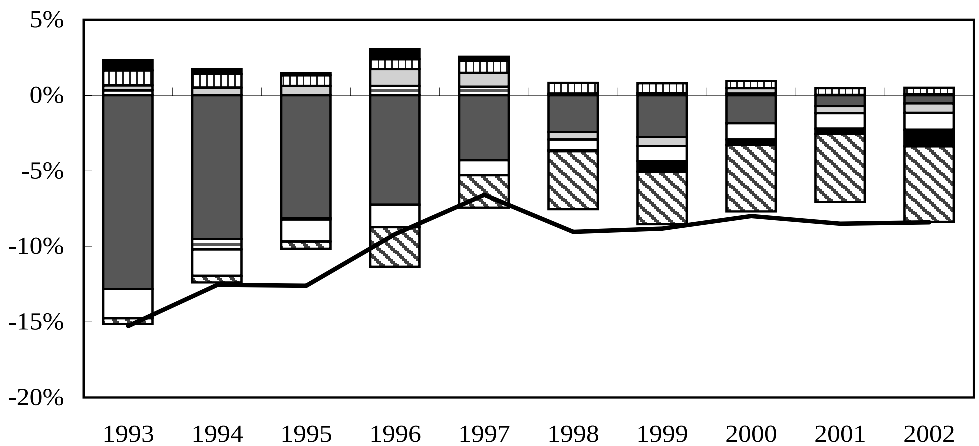
In this paper, based on newly constructed prefectural land price data, we first estimated the long-run equilibrium relationships by means of a panel cointegration analysis. We

Figure 6: Contributions to Changes in Land Prices

(1) Tokyo



(2) Urban other than Tokyo



(3) Non-Urban

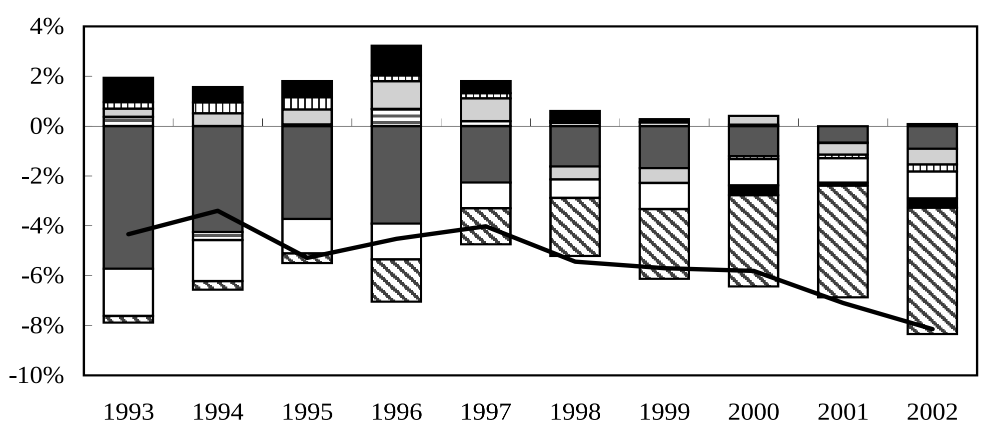
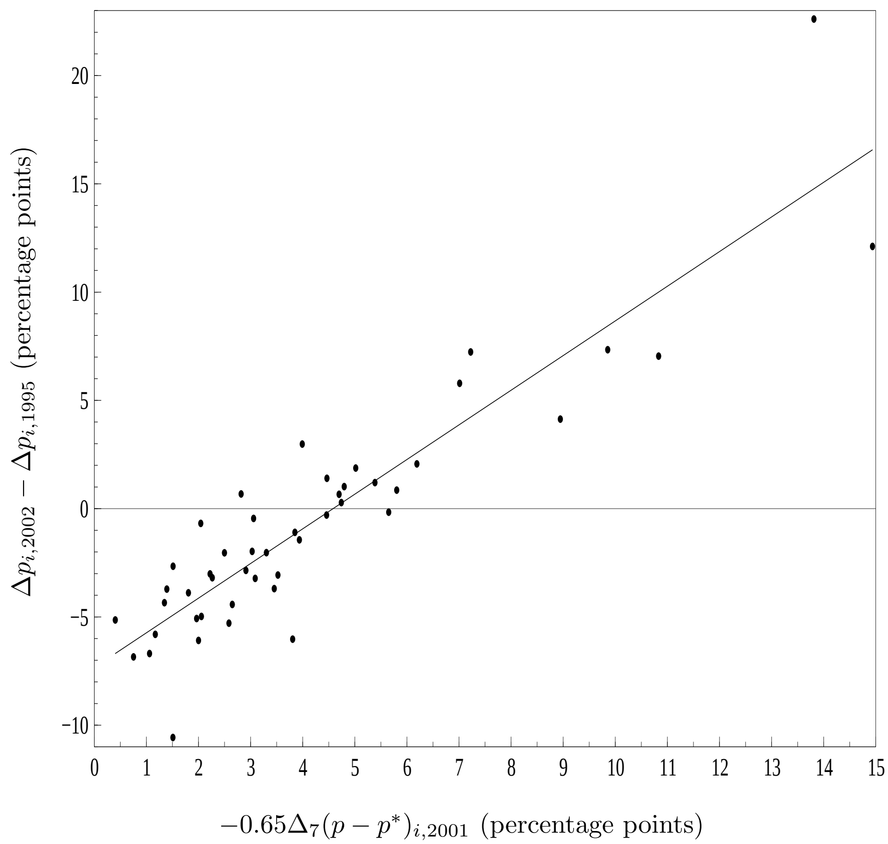


Figure 7: Contributions of $(p - p^*)_{i,t-1}$



then estimated an error-correction model of land prices, which embeds the above long-run equilibrium.

Our main findings are summarized below:

1. The newly constructed chain-type index of prefectural land prices highlights regional discrepancies: in recent years, land prices in Tokyo have bottomed out, but prices in other urban areas albeit at a slower pace, have continued to decline significantly, and prices in non-urban areas have fallen at a faster pace.

2. The panel cointegration analysis reveals that the PVR cum price expectation can be regarded as a long-run equilibrium relationship. However, no cointegration relationship can be found for the fundamental price relation.

3. The ECM reveals that deviations from the long-run equilibrium and non-performing loans have sizable effects on land prices. Moreover, the recent regional discrepancies in land prices are closely related to deviations from the long-run equilibrium.

These findings suggest that land prices in Tokyo have already declined to a level in line with the long-run equilibrium prices calculated from price expectation, nominal income and capital cost, but prices in the other areas have not yet fully adjusted to reach these equilibrium levels.

Since the NPL problems for banks, and the debt-overhang problems for firms, are different sides of the same coin, the significant NPL ratio in the ECM is consistent with the work of Sekine and Tachibana (2004), who find the deterioration in financial conditions has induced firms to sell a vast amount of land in the 1990s.

Given the recent deviations from the long-run equilibrium prices, land prices in Tokyo are likely to increase more clearly, as banks dispose their NPLs; but it is likely to take some time before land prices in other areas, especially in non-urban areas, level off.

Appendix A: Panel Unit-Root Tests

This appendix describes additional results of panel unit-root tests.

We first conduct, by prefecture, Augmented Dicky-Fuller (ADF) tests for level and first difference of land prices p_{it}, nominal income y_{it}, and capital cost r_{it} (Table A.1). The number of prefectures for which the null of unit-root are rejected is as follows:

Variables	p	y	r
Level	1	37	0
First Difference	1	2	47
Sample Periods	1977-2003	1976-2001	1979-2002

For p_{it} and Δp_{it}, only a single prefecture rejects the null hypothesis; i.e., p_{it} is more than I(2). For y_{it} almost all prefectures reject the null, whereas for Δy_{it} only two prefectures reject it; they are inconsistent in that the level variables indicate I(0), but the first differenced variables indicate more than I(2). Finally, for r_{it}, no prefecture rejects the null hypothesis, and for Δr_{it} all prefectures reject the null; the result strongly supports the claim that r_{it} is I(1).

It is well known, however, that these unit-root tests have low power and often fail to reject the null, especially in the case of small samples as here. Panel unit-root tests aim to mitigate low power by adding cross-sectional variations. As well as the Hadri panel unit-root tests described in the main text, this appendix, following Maddala and Wu (1999), sets out the Fisher panel unit-root tests.

For variable x_{it}, the Fisher test can be written as:

$$\lambda = -2 \sum_{i=1}^{N} \ln \pi_i,$$

where π_i are p-values of ρ_i in the ADF tests for each prefecture ($i = 1, ..., N$):

$$\Delta x_{it} = \rho_i x_{i,t-1} + \sum_{j=1}^{p_i} \theta_{ij} \Delta x_{i,t-j} + \alpha_i + \epsilon_{it}. \tag{A.1}$$

In the absence of cross-sectional correlation among the ϵ_{it}, λ follows the χ^2 distribution with $2N$ degrees of freedom under the null hypothesis that "x_{it} for all prefectures has a unit-root," against the alternative hypothesis that "x_{it} does not have a unit-root for at least one prefecture."[8]

[8] The null hypothesis is H_0: $\rho_i = 0$ for all i, while the alternative hypothesis is H_1: $\rho_i < 0$ for at least one i. This alternative hypothesis is same as that for the IPS test (Im, Pesaran, and Shin, 2003). Meanwhile, the alternative hypothesis of the Levin-Lin test (Levin, Lin, and Chu, 2002) is H_1': $\rho_i = \rho < 0$, which requires the common value ρ under the alternative. Less restriction under the null is thought to make both the IPS and the Fisher more appealing in a heterogeneous panel such as our prefectural land prices.

Table A.1: Augmented Dickey-Fuller Tests

Prefectures	p	Δp	y	Δy	r	Δr
Hokkaido	(1) -3.068*	(1) -1.671	(2) -8.058**	(2) -1.358	(0) -1.064	(1) -4.989**
Aomori	(1) -1.648	(0) -0.030	(0) -6.963**	(0) -3.230*	(0) -1.043	(0) -3.824**
Iwate	(1) -1.880	(0) -1.433	(1) -5.579**	(1) -2.277	(0) -1.014	(1) -4.339**
Miyagi	(1) -2.330	(0) -1.194	(0) -2.925	(0) -1.454	(0) -1.158	(1) -4.703**
Akita	(1) -1.692	(0) -0.747	(2) -5.718**	(2) -2.491	(0) -1.067	(0) -3.854**
Yamagata	(1) -1.865	(0) -0.375	(0) -7.218**	(0) -3.222*	(0) -0.815	(0) -3.986**
Fukushima	(3) -2.259	(2) -0.577	(0) -2.767	(0) -2.057	(0) -1.271	(1) -4.925**
Ibaraki	(1) -2.421	(1) -1.323	(0) -5.064**	(0) -2.840	(0) -1.403	(1) -5.082**
Tochigi	(3) -2.150	(2) -1.126	(2) -6.657**	(2) -0.733	(0) -1.132	(1) -5.062**
Gunma	(1) -2.286	(1) -2.143	(0) -7.238**	(0) -2.031	(0) -1.651	(1) -5.644**
Saitama	(1) -1.636	(0) -2.902	(0) -2.414	(0) -1.559	(0) -1.831	(1) -4.861**
Chiba	(1) -1.653	(0) -2.281	(2) -6.715**	(2) -0.621	(0) -1.824	(1) -4.670**
Tokyo	(1) -2.715	(0) -1.540	(0) -2.562	(0) -0.796	(1) -2.251	(0) -3.709*
Kanagawa	(1) -1.631	(0) -3.149*	(0) -6.767**	(0) -1.913	(0) -1.394	(1) -4.373**
Niigata	(1) -2.045	(0) -0.789	(0) -5.578**	(0) -2.654	(0) -1.200	(1) -4.759**
Toyama	(1) -1.882	(0) -1.202	(1) -6.945**	(1) -1.599	(0) -1.116	(1) -4.711**
Ishikawa	(1) -1.899	(0) -1.088	(2) -6.392**	(2) -0.970	(0) -1.735	(1) -5.595**
Fukui	(1) -1.956	(0) -1.176	(1) -6.549**	(1) -1.585	(0) -1.704	(1) -5.327**
Yamanashi	(1) -1.954	(0) -1.670	(0) -5.379**	(0) -2.669	(0) -1.248	(1) -5.535**
Nagano	(3) -1.962	(1) -1.731	(0) -6.018**	(0) -2.287	(0) -1.435	(1) -4.998**
Gifu	(3) -2.347	(2) -0.921	(2) -8.098**	(2) -0.731	(0) -1.334	(1) -5.331**
Shizuoka	(3) -2.111	(1) -2.601	(1) -4.753**	(1) -0.676	(0) -1.580	(1) -5.218**
Aichi	(1) -2.764	(1) -1.920	(2) -6.176**	(2) -1.139	(0) -1.526	(1) -4.861**
Mie	(1) -2.121	(1) -2.215	(1) -3.977**	(1) -1.115	(0) -1.424	(1) -5.388**
Shiga	(3) -2.113	(2) -1.461	(0) -6.095**	(0) -3.020*	(0) -1.459	(1) -4.933**
Kyoto	(2) -1.678	(1) -2.745	(3) -6.725**	(3) -1.119	(0) -1.881	(1) -4.775**
Osaka	(1) -2.907	(1) -1.825	(2) -4.850**	(2) -1.044	(0) -1.389	(1) -4.574**
Hyogo	(1) -2.342	(1) -2.459	(2) -4.954**	(2) -1.310	(0) -1.539	(1) -4.675**
Nara	(3) -2.060	(2) -1.691	(0) -6.545**	(0) -1.959	(0) -1.410	(1) -4.913**
Wakayama	(3) -1.839	(2) -1.607	(0) -3.547*	(0) -4.073**	(0) -1.366	(1) -4.892**
Tottori	(1) -1.996	(0) -2.125	(1) -9.152**	(1) -1.922	(0) -1.541	(1) -5.092**
Shimane	(3) -2.335	(3) 0.290	(0) -4.591**	(0) -3.989**	(0) -0.732	(0) -3.902**
Okayama	(1) -1.924	(0) -1.917	(0) -3.139*	(0) -1.565	(0) -0.675	(1) -4.925**
Hiroshima	(2) -1.675	(1) -2.055	(0) -2.734	(0) -1.354	(0) -0.801	(1) -4.820**
Yamaguchi	(1) -2.095	(1) -1.026	(0) -5.187**	(0) -2.789	(0) -0.802	(1) -4.393**
Tokushima	(3) -1.861	(2) -0.853	(1) -5.901**	(1) -1.340	(0) -1.084	(1) -5.012**
Kagawa	(1) -1.430	(0) -2.395	(0) -2.958	(0) -2.270	(0) -0.988	(1) -5.029**
Ehime	(1) -1.779	(2) -1.061	(2) -4.641**	(2) -2.322	(0) -1.145	(1) -4.988**
Kochi	(1) -1.260	(0) -2.534	(0) -2.382	(0) -2.020	(0) -1.177	(1) -4.561**
Fukuoka	(1) -2.184	(0) -1.352	(0) -3.253*	(0) -1.613	(0) -1.113	(1) -5.381**
Saga	(1) -1.959	(0) -1.637	(2) -7.774**	(2) -2.593	(0) -0.883	(1) -4.499**
Nagasaki	(1) -1.866	(0) -1.404	(0) -5.506**	(0) -2.095	(0) -1.149	(1) -5.360**
Kumamoto	(1) -2.235	(1) -1.621	(0) -7.820**	(0) -2.191	(0) -1.084	(1) -5.108**
Oita	(1) -2.180	(2) -1.094	(0) -5.946**	(0) -2.298	(0) -0.893	(1) -4.553**
Miyazaki	(3) -2.717	(3) -1.795	(3) -6.830**	(3) -1.466	(0) -0.953	(0) -4.114**
Kagoshima	(1) -1.785	(0) -1.603	(3) -8.939**	(3) -2.641	(0) -1.425	(1) -4.769**
Okinawa	(1) -2.001	(0) -0.828	(0) -7.958**	(0) -2.212	(0) -2.021	(1) -5.239**

Note: ADF t-values. Numbers in parentheses are the lag lengths, which are determined by the 10% significance level. "**" and "*" denote statistical significance at the 1% and 5% levels, respectively.

Table A.2: Panel Unit-Root Tests (Fisher)

	Level			First Difference		
	Fisher	1% CV	5% CV	Fisher	1% CV	5% CV
p	281.81*	315.40	228.25	203.61*	238.91	190.04
y	1027.90**	238.88	189.98	165.77*	198.14	158.96
r	150.10	249.72	200.98	860.97**	366.96	257.04

Note: "**" and "*" denote statistical significance at the 1% and 5% levels, respectively. The corresponding critical values (CV) are obtained from bootstrap simulations of 10,000 replications.

If the absence of cross-sectional correlation among ϵ_{it} is suspicious, Maddala and Wu (1999) recommend obtaining critical values by bootstrap instead of using the χ^2 distribution. This appears to be the case for our prefectural land prices, in view of anecdotes that the property price bubble in Tokyo has infected other urban areas and then non-urban areas. We therefore conduct bootstraps to obtain the relevant critical values.[9]

Table A.2 summarizes the test results: p_{it} and y_{it} are I(0) since, for these variables, the null of unit-root is rejected both in level and in first difference. But r_{it} is I(1), because the null is not rejected in level, but is rejected in first difference.

It is strange that stationarity is observed for land price and nominal income, for which many time-series analyses find nonstationarity—the Hadri tests in the main text also indicate that these variables are I(1).

We suggest that the stationarity is due to one or more of the following statistical reasons, and assume tentatively that these variables are indeed I(1).

- Distortion in size: According to the Monte Carlo experiments by Maddala and Wu (1999), the power and size of the Fisher test are more precise than for the Levin-Lin

[9]The bootstrap is conducted according to the procedure below:

1. Calculate λ by estimating equation (A.1) by prefecture (the lag lengths are determined by the 10% significance level).
2. Calculate ϵ^0_{it} by estimating $\Delta x_{it} = \rho^0_i \Delta x_{i,t-1} + \epsilon^0_{it}$ for each prefecture (the S_3 sampling scheme in Maddala and Kim (1999)).
3. Preserving the cross-sectional correlation of ϵ^0_{it}, obtain the bootstrap sample of ϵ^*_{it} by randomly resampling ϵ^0_{it} with replacement.
4. Construct the series x^*_{it} from $x^*_{it} = x^*_{i,t-1} + \rho^0_i \Delta x^*_{i,t-1} + \epsilon^*_{it}$. x^*_{i0}. Here x^*_{i1} are obtained by randomly resampling a block of x_{it} with replacement.
5. Calculate λ from x^*_{it} by Step 1.
6. Repeat Steps 3-5, 10,000 times.

Sort the obtained λ and assign the 1% critical value as the 100th largest λ (the 5% critical value as the 500th largest λ).

test and the IPS test. However, they also report a distortion in size with short time-series data such as 25 periods. This is the case for our data set, so that the rejection of the null for p_{it} and y_{it} might be due to the distortion in size.

- Deterministic trend: We do not include deterministic trends, because in most prefectures they are not significant in the ADF tests; for p_{it} and y_{it} respectively, deterministic trends are significant at the 10% (20%) level, in 14 (21) and 17 (27) prefectures. However, these appear to be non-negligible numbers. By adding deterministic trends to equation (A.1) as:

$$\Delta x_{it} = \rho_i x_{i,t-1} + \sum_{j=1}^{p_i} \theta_{ij} \Delta x_{i,t-j} + \alpha_i + \delta_i t + \epsilon_{it},$$

  and then conducting the Fisher tests with deterministic trends, λ becomes 95.16 for p_{it} and 82.99 for y_{it}; the null hypotheses are not rejected at the 5% significance level (the corresponding critical values are 164.09 and 130.63 respectively).

- Bootstrap: The above bootstrap takes into account cross-sectional correlations among error terms, but assumes no auto-correlations nor heteroscedasticity. As with Herwartz and Reimers (2002), who use the wild bootstrap, it might possible to obtain better results by taking into account these complications.

Appendix B: Effective Tax Rates on Real Properties

Taxes on real properties in Japan are categorized into three groups according to when these taxes are levied: acquisition, holding and sales (Yamazaki, 1999; Kanemoto, 1990).

In calculating effective tax rates on real properties, we take into account various taxes on land acquisition (Real Property Acquisition Tax and Registration and License Tax) and on land holding (Property Tax, Urban Planning Tax and Land Value Tax). We do not take into account Special Landholding Tax, because this tax is applied only to the properties in prescribed districts (partially urbanized areas—*shigaika chosei kuiki*—in certain municipalities) in the Tokyo, Osaka and Nagoya metropolitan areas. Nor do we take into account Enterprise Tax, since it is levied on acquisitions and holdings of business buildings, but not on land itself.

We omit taxes on land sales in the following calculation, because of their notorious complexity. In addition to various exemptions, the tax rates on capital gains from property sales differ according to (i) how long the properties have been owned: "long-term", "short-term" and "excessively short-term" (definitions of these terms have altered from time to time) and (ii) who has owned the properties: depending on this, either Individual Income Tax rates (which have progressive structure) or Corporate Income Tax rates are applied. Furthermore, Inheritance Tax and Gift Tax increase the complexity of the calculation—see, for instance, Nishimura et al. (1999) on the preferential treatments of farmers in housing areas with respect to Inheritance Tax.

Effective tax rates of each property tax are calculated as:

$$\text{Effective Tax Rates} = \text{Evaluation Ratios} \times \text{Statutory Tax Rates},$$

where the evaluation ratios are obtained as follows:

- For Property Tax, following Yamazaki and Idee (1997), we derive the evaluation ratios for each prefecture as the ratio of its effective tax bases (MPHPT, "Summary Report on Prices of Fixed Assets") to values of land assets in the SNA statistics.
- For Real Property Acquisition Tax, Registration and License Tax, and Urban Planning Tax, we multiply the evaluation ratios of Property Tax by the mark-up rates tabulated in Table B.1.
- For Land Value Tax, we set its evaluation ratio 0.8.

These individual effective tax rates add up to effective tax rates on real properties τ_{it}.

Table B.1: Statutory Tax Rates (%)

	PT	RPAT	RLT	UPT	LVT
1975	1.4	4.0 (1.00)	5.0 (1.00)	0.2 (1.00)	...
1976	1.4	4.0 (1.00)	5.0 (1.00)	0.2 (1.00)	...
1977	1.4	4.0 (1.00)	5.0 (1.00)	0.2 (1.00)	...
1978	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1979	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1980	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1981	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1982	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1983	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1984	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1985	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1986	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1987	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1988	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1989	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1990	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1991	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	...
1992	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	0.20 [0.80]
1993	1.4	4.0 (1.00)	5.0 (1.00)	0.3 (1.00)	0.30 [0.80]
1994	1.4	4.0 (0.50)	5.0 (0.40)	0.3 (1.00)	0.30 [0.80]
1995	1.4	4.0 (0.67)	5.0 (0.40)	0.3 (1.00)	0.30 [0.80]
1996	1.4	4.0 (0.50)	5.0 (0.40)	0.3 (1.00)	0.15 [0.80]
1997	1.4	4.0 (0.50)	5.0 (0.40)	0.3 (1.00)	0.15 [0.80]
1998	1.4	4.0 (0.50)	5.0 (0.40)	0.3 (1.00)	0.00 [0.00]
1999	1.4	4.0 (0.50)	5.0 (0.33)	0.3 (1.00)	0.00 [0.00]
2000	1.4	4.0 (0.50)	5.0 (0.33)	0.3 (1.00)	0.00 [0.00]
2001	1.4	4.0 (0.50)	5.0 (0.33)	0.3 (1.00)	0.00 [0.00]
2002	1.4	4.0 (0.50)	5.0 (0.33)	0.3 (1.00)	0.00 [0.00]

Notes:

1. PT: Property Tax, RPAT: Real Property Acquisition Tax, RLT: Registration and License Tax UPT: Urban Planning Tax, LVT: Land Value Tax.
2. Figures in parentheses are mark-up rates to the evaluation ratios of Property Tax. Figures in squared brackets are evaluation ratios.
3. Land Value Tax was introduced in 1992, and its imposition was temporarily suspended from 1998.

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