Empirical Modelling of Regional House Prices and the Ripple Effect

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Abstract

This paper investigates the stationarity properties and long-run relationship of regional house prices in Taiwan during 1993Q1 to 2009Q2. We apply the recent unit root test of the panel seemingly unrelated regressions augmented Dickey-Fuller (SURADF) test developed by Breuer et al. (2001). Our empirical results illustrate that regional house prices in Taiwan are a mixture of stationary and non-stationary processes, and that the traditional panel unit root tests could lead to misleading inferences. Second, for the estimated half-lives in Taipei City, the degrees of mean reversion are greater than those for the overall Taiwan area. Third, our findings provide substantive evidence in favor of the existence of a long-run equilibrium relationship among Taipei County, Taoyuan-Hsinchu, Taichung, and Tainan-Kaohsiung. Finally, the results of the weak exogeneity test indicate that uni-directional causality relationships or ripple effects exist for three regions - Taipei County, Taoyuan-Hsinchu, and Tainan-Kaohsiung - to Taichung.

Keywords: Regional house prices; Ripple effect; Cointegration; Causality; Panel SURADF test.

1. Introduction

The issue of regional house prices has shown increasing interest and investigation over the past several years both from researchers and government policy makers. This is principally through the marked changes in asset prices, which have had considerable influence on housing affordability. The Taiwan context provides a substantial challenge for empirical models of regional house prices, because geographical and regional economic conditions differ much in Taiwan, which are reflected in the value of housing prices.

According to the census of Taiwan's Ministry of Interior in 2006, the average home ownership rate is over 87%, which is the highest in the world. Because there are few investment instruments and the relevant opportunity cost of housing is better than other assets in Taiwan, thus the housing market offers a beneficial investment opportunity (Chen et al., 2007). To date there are very few empirical studies that have investigated the interrelationship between regional house prices for Taiwan. We employ regional data controls for such national characteristics and focus on specific aspects of the economic and financial systems. We intend to see whether the housing market performance of Taiwan can be ascribed. As to the housing policy in Taiwan, there is a much smaller public housing sector in Taiwan, and thus private housing developers have more freedom and face less competition from the state sector. More importantly, while home prices are more responsive to regional economic and demographic shocks instead of national shocks, a highlight at the regional level enables us to compare the dynamics in housing markets across regions (Kim and Bhattacharya, 2009).

The behaviour of regional house prices has established a bright area of research in recent years (Holmes and Grimes, 2008). As to the empirical studies of Taiwan's house prices, little attention is paid to examining the diffusion of regional house prices in different areas. Some studies have investigated housing prices and quality (Huang, 1999), some have examined the effects of economic variables on housing prices (Hsueh and Chen, 1998), and still others have applied the structural time-series model to examine housing prices (Chen, 2003). All of these papers except Chien (2010) did not investigate the spatial diffusions among Taiwan's regional house prices. However, Chien only focuses on the ripple effect of the regional/national house price ratios, but not the long-run equilibrium and lead-lag relationship among regional house prices in different areas.

Some studies in the literature have focused on housing dynamics through different theoretical approaches, including neighborhood change, filtering, search, equity effects, and urban growth. Empirical investigations of these models are uncommon due to the complexity of the models or being short of data. Some have studied the interdependence between housing and financial assets, such as Jud and Winkler (2002) and Riddel (2004). Another important line of empirical research refers to whether the disparities in housing prices occur irregularly or whether a ripple effect exists within regional house markets in the long run. This subject is always discussed as a convergence or divergence in regional house prices in the existing literature. Alexander and Barrow (1994) state that regional house prices are not considered by economic theory to have a common trend over time, but rather the migration of households for economic changes within regions causes the possibility of convergence in regional house prices. Meen (1999) also indicates that convergence exists if long-run equilibrium relationships occur between the regional house markets.

Because housing is an immobile asset, property can cause regional house price variations to be persistent or show non-stationarity (Ashworth and Parker, 1997). To seek explanations of the structure of regional house prices within the UK, some works apply Engle and Granger (1987) or Johansen (1988) cointegration tests to investigate the notion of a causal link between different regional prices for houses, but the conclusions drawn from these relative studies are diverse. Using standard house price models on regional data, MacDonald and Taylor (1993) and Alexander and Barrow (1994) demonstrate a cointegration relationship between regional house prices, but within only either the North or the South of Britain. This finding has prompted a weak segmentation of North/South housing markets in the UK. Conversely, Ashworth and Parker (1997) cast doubt that different UK regions appear to adjust to housing price shocks together using the error correction model (ECM).

If a ripple effect or convergence is present, then the ratio between each regional price and the national house price is stationary (see Meen, 1999; Cook, 2005; Holmes and Grimes, 2008). Employing the augmented Dickey and Fuller (ADF; Dickey and Fuller, 1979) unit root test, the empirical results of Meen (1999) do not support stationarity in the regional-national house price ratios for the UK. After revising the model with spatial correlations in housing prices, Meen (1999) achieves the reverse conclusion. Using the threshold regression model and allowing for the possibility of an asymmetric adjustment process about the stationary attractor, Cook (2003) supports the stationarity of regional house price ratios in the UK. Cook (2005) applies a joint application of two unit root tests, the DF-GLS test of Elliot et al. (1996), and the KPSS test of Kwiatkowski et al. (1992), and the empirical results confirm the existence of stationarity in regional house price ratios in the UK. Being different from past research by examining the issue of whether regime changes break down the stability of the ripple effect, Chien (2010) applies the Lagrange multiplier unit root test to support the existence of a ripple effect for each city in Taiwan except Taipei City.

Few research studies investigate the spillover of housing price changes within neighboring areas. Clapp et al. (1995) find evidence of spatial diffusion for housing price changes between neighboring towns in Connecticut and near San Francisco, but there is none across non-neighboring towns. The empirical results of Dolde and Tirtiroglu (1997) also support the same results by GARCH-M models. Others investigate the ripple effect of a housing submarket within a city. Ho et al. (2008) examine spatial ripple effects across different quality tiers of housing within Hong Kong for the period 1987 to 2004.

The main purpose of this paper is to investigate the stationarity properties and long-run relationship of regional house prices in Taiwan with quarterly data over 1993Q1 to 2009Q2. To examine the stationarity of regional house prices, we apply the newest panel seemingly unrelated regressions augmented Dickey-Fuller (Panel SURADF) test developed by Breuer et al. (2001), which allows us to account for possible cross-sectional effects and to identify how many and which regional house prices within the panel contain a unit root (non-stationary). We proceed by measuring the half-lives and the corresponding confidence intervals when stationarity of the regional house prices is confirmed. Third, once we confirm the stationary behaviour for the regional house price series, we then investigate the relationships between the series based on the different order of integration. For regional house prices in which the series are integrated of degree one (I(1)), we use the cointegration method to evaluate whether these regional house prices are cointegrated or segmentated, indicating market integration (convergence) or segmentation (divergence) in these regional house prices (MacDonald and Taylor, 1993). Finally, we implement the weak exogeneity test to investigate the causality relationships and examine the ripple effects among different regional house prices.

The remainder of the paper proceeds as follows. Section 2 outlines and discusses Taiwan's housing market and regional economic development. Section 3 describes the methodology. Section 4 presents the empirical findings, and Section 5 summarizes the conclusions that are drawn.

2. Taiwan's Housing Market and Regional Economic Development

Real estate is enormously important in Taiwan due to people's belief in the traditional idea of 'land is wealth'. According to the census of Taiwan's Ministry of Interior in 2006, the average home ownership rate is over 87%, which is the highest in the world. Moreover, the average housing unit vacancy rate is 17.6%, which is far above the average of 3-5% in other countries. Taiwan's space consumption hit $42m^2$ per person in 2008. In terms of living space, Taiwan surpassed many advanced countries, for instance, UK, Singapore, and Japan (Yip and Chang, 2003; Population Census, 2000).

With a population of 23 million and an area of 36,000 square kilometers, Taiwan is one of the most densely populated areas in East Asia. Taiwan's population migration to urban areas has been considerable since the 1970s. The main metropolitan areas are in Figure 1 and are introduced as follows. First, the northern region includes Taipei City, being the capital and most important economic center, and its two surrounding areas of Taipei County and Taoyuan-Hsinchu. In 2009 the residents in Taipei City amounted to 2.6 million, and there are respectively 3.9 million in Taipei County and 2.5 million in Taoyuan-Hsinchu. Next, Taichung is in the central region and Tainan-Kaohsiung is in the southern region, which respectively contain around 2.6 and 3.8 million residents.

The expansion of these areas in Taiwan has been subject to economic development and changing global circumstances. In the 1960s and 1970s, Taiwan's spatial economy was characterized by a bipolar concentration, with one core of the southern area being Kaohsiung, with Taiwan's largest seaport home to heavy and petrochemical industries, and the core of northern area being the capital, Taipei, which has became a highly diversified regional economy with strong political and corporate power (Lin and Liaw, 2000). At the same time, a smaller Export Processing Zone (EPZ) near Taichung was established in 1969, which led to the industrialization of that city. Although Taichung and Kaohsiung took advantage of fast industrialization starting in the 1960s, the traditional manufacturing industries in these two areas began to lose their competitiveness in the 1990s. Conversely, at the same time Taipei City upgraded itself into the node of the high-technology knowledge center and has the status of a regional global city (Wang, 2003). In Taipei there are many new jobs which are created from many corporate headquarters and advanced service industries.

Geographical and regional economic conditions differ much in Taiwan, which is reflected in the value of housing prices. Table 1 shows the average house price in these areas. In the fourth quarter of 2008, Taipei City's average house price is NT$9,600,000 and is at least 50% higher than other areas, reflecting the relative economic gains of Taipei City, coupled with an inelastic supply of housing. The average house price of Kaohsiung is NT$4,940,000, the lowest in all regions, which also shows the loss of economic competition in Kaohsiung being attributed to the shocks of economic restructuring and globalization in the 1980s.

As to the housing policy in Taiwan, there is a much smaller public housing sector, providing rental and owner-occupied housing to approximately 8% of all households only (Chiu, 2010). Because of the housing policy heavily skews towards home ownership, applying subsidized mortgage loans was an alternative to direct provision and started in 1990. As the level of housing intervention is relatively low in Taiwan, private housing developers have more freedom and face less competition from the state sector[1] (Yip and Chang, 2003). In other words, the government of Taiwan acts with a much smaller role in the housing sector, which makes the housing market operate on free market principles.

Another special feature of Taiwan's housing market is the pre-sale system. Having experienced rapid economic development in Taiwan, the pre-sale system has been induced for the need of establishing more efficient markets where private property can be created. Developers sell their property before building is started to get financing for their businesses and to reduce the risk of building property that can be empty. One characteristic feature of the pre-sale system is its similarity to a forward or futures deal (Chang and Ward, 1993), which can improve the efficiency of the housing market.

3. Methodology

3.1 Panel unit root tests

Recent developments in panel unit root testing include refinements made to the Levin et al. (2002) test (LLC), the Breitung (2000) test, the Im et al. (2003) test (IPS), the Fisher-type ADF, the Phillips-Perron test (see Choi, 2001 and Maddala and Wu, 1999), and the Hadri (2000) test. The LLC test is based on the ADF test, but in a panel setting the model is expressed as follows:

$$\Delta y_{it} = \alpha_i + \beta_i y_{i,t-1} + \sum_{j=1}^{p_i} \rho_{ij} \Delta y_{i,t-j} + e_{it} \,, \tag{1}$$

where β_i is restricted such that it is identical across regions; $y_{it}\ (i = 1,2,\ldots,N; t = 1,2,\ldots,T)$ is the house price series of panel member regions *i* in period *t*; p_i is the number of lags in the ADF regression; and the error terms e_{it} are assumed to be independently and normally distributed random variables for all *i*'s and

[1] Yip and Chang (2003) indicate that "Unlike Hong Kong, where state monopolizes land supply, and also in Singapore, where the powerful Land Acquisition Act enables the state to control land, land in Taiwan is mostly privately owned. Contrary to private housing development in South Korea, private developers in Taiwan enjoy supplementary credits from the state for bailing them out during a market slump."

t's with zero means and finite unit specific variances, σ_i^2.

The IPS (2003) test relaxes the assumptions of the LLC (2002) test by allowing β of equation (1) to vary across units under the alternative hypothesis. It is also more general in the sense that it allows for heterogeneity in the autoregressive coefficients for all panel members. The alternative hypothesis simply implies that some or all of the individual series are stationary. However, Breitung (2000) shows that the IPS (2003) test can suffer from a loss of power due to bias correction when individual specific trends are included. He proposes an alternative test without bias adjustments and shows that it has higher power than the IPS test. The Breitung (2000) test statistic examines the null hypothesis that the panel series is difference stationary, and the alternative hypothesis assumes that the panel series is stationary.

In order to have a more powerful test, Hadri (2000) argues that one should have stationarity under the null hypothesis, and it should be reversed in order to have a stronger power test. An alternative approach to the panel unit root tests uses Fisher's (1933) results to derive tests that combine the p-values from individual unit root tests. Maddala and Wu (1999) offer their alternative test based on the p-values of the separate Dickey–Fuller unit root tests for each of the N cross-section units. The Maddala and Wu (1999) test has the advantage over the Im et al. (2003) test in that it does not depend on different lag lengths in the individual ADF regression. An important advantage of this test is that it can be used regardless of whether the null is one of integration or stationarity.

3.2 The panel SURADF unit-root tests

Breuer et al. (2001, 2002) develop a panel unit root test that is based on the augmented Dickey-Fuller (ADF) regression estimation in a seemingly unrelated regressions (SUR) framework and then test for an individual unit root within the panel members. This procedure has several advantages. First, these multivariate tests use the information content in the variance–covariance matrix, so that the unrealistic assumption of cross-section independence made in the panel tests can be avoided. Second, conventional univariate unit root tests not only fail to consider information across regions, but are also restricted regarding the problem of a small sample, thereby leading to less efficient estimations (see Chiu, 2002). Exploiting the information from the error covariance and allowing for an autoregressive process will produce more efficient estimators than the single-equation methods. Third, the estimation tests also allow for an important degree of heterogeneity in the lag structure across the panel members, in that the lag order of the augmented test varies among the individuals and the autoregressive parameter also differs for every cross-section. Fourth, the panel SURADF unit root test allows us to identify how many and which members of the panel contain a unit root. However, a common feature of the panel tests is that they maintain the null hypothesis of a unit root in all panel members. Therefore, their rejection indicates that at least one panel member is stationary, although there is no information regarding which ones are stationary.

The unit root test of the panel SURADF for N regions and T time periods is based on the system of ADF equations which can be represented as:

$$
\begin{aligned}
\Delta Z_{1,t} &= \alpha_1 + \beta_1 Z_{1,t-1} + \eta t + \sum_{j=1}^{k1} \varphi_{1,j} \Delta Z_{1,t-j} + e_{1,t} \qquad t = 1,2,....,T \\
\Delta Z_{2,t} &= \alpha_2 + \beta_2 Z_{2,t-1} + \eta t + \sum_{j=1}^{k2} \varphi_{2,j} \Delta Z_{2,t-j} + e_{2,t} \qquad t = 1,2,....,T \\
&\vdots \qquad\qquad\qquad\qquad\qquad \vdots \\
\Delta Z_{N,t} &= \alpha_N + \beta_N Z_{N,t-1} + \eta t + \sum_{j=1}^{kN} \varphi_{N,j} Z_{N,t-j} + e_{N,t} \qquad t = 1,2,....,T \ , \qquad (2)
\end{aligned}
$$

where Z denotes real regional house prices for the overall Taiwan area (LT), Taipei City (LTC), Taipei County (LTCY), Taoyuan-Hsinchu (LTH), Taichung (LTA), and Tainan-Kaoshiung (LTK), respectively, and $e_{i,t}$ $(i=1,2,....,N)$ is an error term. Coefficient α_i is the heterogeneous constant term, $\beta_i=\rho_i-1$, and ρ_i is the autoregressive coefficient for the *i*th cross-sectional member of the series, while *t* denotes the deterministic time trend.

Equation (2) tests the null hypothesis of a unit root against the trend stationarity. This system is estimated by the SUR procedure, and we test the N null (H_o^i) and alternative hypotheses (H_A^i) individually as:

$$
\begin{aligned}
&H_0^1:\beta_1=0; H_A^1:\beta_1<0\\
&H_0^2:\beta_2=0; H_A^2:\beta_2<0\\
&\vdots \qquad\qquad \vdots\\
&H_0^N:\beta_N=0; H_A^N:\beta_N<0,
\end{aligned}
\tag{3}
$$

with the test statistics being computed from the SUR estimates of system (2), while the critical values are generated by Monte Carlo simulations.

3.3 The analysis of half-lives

Even more importantly, the unit root tests are uninformative as to the speed of mean reversion. Alternatively, the “half-life” of deviation - which is defined as the number of periods required for a unit shock to dissipate by one half - measures the degree of mean reversion and the speed of adjustment back towards the long-run equilibrium. To motivate this measure, suppose that the deviations of the real regional house prices $Z_{i,t}$

from the long-run value $Z_{i,0}$ follow an AR(1) process:

$$Z_{i,t} - Z_{i,0} = \alpha\left(Z_{i,t-1} - Z_{i,0}\right) + \varepsilon_{i,t}\,, \tag{4}$$

where ε is a white noise. The half-life deviation h is defined as the horizon at which the percentage deviation from the long-run equilibrium is one half - that is:

$$\alpha^{h} = \frac{1}{2} \quad \Rightarrow \quad h = \frac{\ln\left(1/2\right)}{\ln(\alpha)}. \tag{5}$$

A conventional 95% confidence interval associated with the above half-life statistic based on normal distributions. Since h cannot be negative, we impose a lower bound of zero (see Rossi (2005) for more details).

4. Empirical Results

4.1 Data description

This empirical work employs six real house price indices of Taiwan, including the national house price index for the overall Taiwan area (LT) and five regional house price indices: Taipei City (LTC), Taipei County (LTCY), Taoyuan-Hsinchu (LTH), Taichung (LTA), and Tainan-Kaohsiung (LTK). This was done over the period 1993Q1 to 2009Q2, where 2001 is the base year. The data are drawn from the housing index database of Cathay Real Estate Development Company. The index represents the price path of pre-sale housing and new houses. Its trait is that it integrates the average price, character, scale, and negotiated price of each single case, and adopts the hedonic housing price model and Laspeyres index formula to fix the character of the housing price and measure the index. The hedonic model is frequently applied to quantify the effect of different housing and neighborhood characteristics on housing prices (Goodman and Thibodeau, 1995). An examination of the individual data series makes it very clear that we require the logarithmic transformations to achieve stationarity in the variance and therefore we transform all of the data series to their logarithmic form.

Figure 2 includes the time series plots of the six house prices included in our model. These figures appear to be non-stationary and exhibit similar patterns. These series show a little variation over the period before 1998, which seem to be breaking around 2000. It is apparent that all of the variables show rising trends, suggesting that, at least initially, we need to include a linear trend in our model. Figure 1 clearly shows that while initially very dispersed, the series exhibits a gradual narrowing of cross regional differences.

4.2 The results of unit-root test and half-lives

When seeking explanations of the structure of regional house prices within Taiwan, if anyone tries to obtain meaningful regression results from a regression containing integrated variables, then it is necessary for these variables to be cointegrated. Therefore, the first step in the analysis is to test for a unit root type of non-stationarity and, if this is confirmed, then a cointegration analysis is the next step.

We start by testing for the presence of a unit root in regional house prices using the ADF (Dickey and Fuller, 1979), DF-GLS (Elliott et al., 1996), PP (Phillips and Perron, 1988), KPSS (Kwiatkowski et al., 1992) and NP (Ng and Perron, 2001) unit root tests.[2] The estimation method adopted in this research utilizes not only the modified Akaike information criterion (MAIC) put forth by Ng and Perron (2001, in the ADF, DF-GLS, and the NP tests for the selection of the optimal lag length, but also the kernel-based criteria put forth by Newey and West (1994) in the PP and the KPSS tests for the selection of the bandwidth. Table 2 reports the results of these univariate unit root tests with intercept and trend. The results show that all variables are non-stationary at the 5%

[2] The null of the KPSS unit-root test is I(0), while the null of the remaining four tests is I(1).

significance level.

In order to provide an analysis of sensitivity and robustness, this paper employs a broad array of panel unit root tests: the LCC (2002), Breitung's (2000) t-statistic, the IPS (2003) W-statistic, the ADF-Fisher Chi-square, the PP-Fisher Chi-square, and Hadri's (2000) Heteroscedastic Consistent Z-statistic. All panel unit root tests are the null hypothesis of non-stationarity to be tested, except for the Hadri (2000) and the Heteroscedastic Consistent Z-statistics test statistic tests with the null hypothesis stationarity. As tabulated in Table 3, all variables are non-stationary at the 5% significance level. We find that it is reasonable to assume all variables follow I(1) processes.

To provide the number of unit roots or stationary cross-section elements, we utilize the Breuer et al. (2001, 2002) test that individually examines the null of non-stationarity in the SUR grid. Table 4 provides the Panel SURADF tests and the critical values for different levels of regional house prices. As the SURADF test has non-standard distributions, the critical values need to be obtained via simulations. In the data generation phase of the simulation, the intercepts and the coefficients on the lagged values for each series are set equal to zero. The estimated 1%, 5%, and 10% critical values are obtained from Monte Carlo simulations based on 66 observations for each series and 10,000 replications by using the lag and covariance structure from the panel of regional house prices. Since the SUR estimation takes into account error correlation, which is different for different data series, the critical values for the SURADF are different for each series.

Our results reveal that these regional house prices are a mixture of I(0) and I(1) processes. As seen in Table 4, we are able to reject the null hypothesis of non-stationarity at the 10% level for the overall Taiwan area (LT) and Taipei City (LTC), showing that housing prices in Taipei City and the Taiwan area are trend stationary (mean-reverting). These series should then return to their trend path over time, and it should be possible to forecast future movements in these two house prices based on past behavior. By contrast, the house price shocks stemming from the other four regions in Taiwan eventually have a permanent effect on regional housing markets.

It is also widely agreed in empirical studies that macroeconomic series are typically affected by the effects of exogenous shocks. In order to obtain robust test results, this study goes further to perform the Zivot and Andrews (1992; ZA hereafter) unit root test with a panel SUR framework. There are two important differences in our testing approach. First, ZA extend the ADF approach by endogenizing the breakpoint determination. Second, the method of SURADF has the capability to consider a structural break by applying ZA estimations, which is called the panel SURZA test. Table 5 provides the estimated breakpoint for each region, the Panel SURZA tests, and the critical values for different regions. The results are almost the same as prior results of the panel SURADF test except for Taipei City. Therefore, after considering the influence of a structure break, we conclude that our important findings - of whether real house prices being stationary or not are affected by regions - do not change.

The unit root test alone may not be sufficient to justify the adjustment dynamics of a long-run equilibrium for regional house prices. It is likely that although the unit-root hypothesis is rejected, deviations are still persistent. What we are interested in is the variable of the speed of convergence to the long-run equilibrium. One measure of the degree of mean reversion that has attracted much attention in the literature is the half-life. The point estimates of the half-lives alone provide an incomplete description of the speed of convergence towards the long-run equilibrium, but to this end, the corresponding confidence intervals are computed to provide better indications of uncertainty around the estimates of half-lives. As can be seen in Table 6 which presents the half-lives and their confidence intervals, the half-lives of two house price indices, Taipei City and the Taiwan area, approximately range from 8.20 to 9.92 quarters. This shows that for Taipei City, the degrees of mean reversion are greater than those for the overall Taiwan area.

4.3. Test for Johansen's multivariate cointegration procedures

According to the empirical results of the Panel SURADF tests, there are four regional house prices that have I(1) processes. Therefore, we further use Johansen's (1988) multivariate maximum likelihood cointegration test to investigate the long-run relationship among the four regional house prices, LTA, LTCY, LTH, and LTK, along with Johansen and Juselius' (1990) cointegrated vector coefficient significance test. Before applying Johansen's (1988) multivariate maximum likelihood cointegration test, we first need to select the appropriate number of lag lengths which should be high enough to ensure that the errors are approximately white noise, but small enough to allow estimations to be made. For this reason, when selecting the number of lag periods, it is necessary that we perform and pass residual misspecification tests using the VAR model. In this respect, numerous economic studies have employed many different lag length selection criteria to determine the autoregressive (AR) lag lengths of time series variables.

An AR process of lag length p refers to a time series in which its current value is dependent on its first p lag value and is normally denoted by AR (p). Note that the AR lag length p is always unknown, and therefore it has to be estimated using various lag length selection criteria, such as the sequential modified likelihood ratio (LR) test, the final prediction error (FPE) method (Akaike 1969), Akaike's information criterion (AIC) (Akaike, 1974), Schwarz information criterion (SC) (Schwarz 1978), and

Hannan-Quinn criterion (HQ) (Hannan and Quinn 1979). All of these criteria are discussed in Lűtkepohl (1991).

We employ the FPE, AIC, SC, and HQ to determine the number of lags that should be used in the VAR, but in the case of a serial correlation we introduce a sufficient number of lags to eliminate a serial correlation of the residuals. From the test results reported in Table 7, we select lag 1. Using this lag length, we perform the test for normality and for the absence of serial correlation in the residuals in the VAR in order to make sure that none of them violates the standard assumptions of the model. To this end, we use the Portmanteau autocorrelation test which computes the multivariate Ljung-Box (1978) Q-statistics for a residual serial correlation up to the specified order.

Under the null hypothesis of no serial correlation up to lag *h*, the statistic is approximately distributed χ^2 with degrees of freedom $k^2(h-p)$, where *p* is the VAR lag order and *k* is the number of endogenous variables. The results indicate that with lag 1, the residuals in the VAR are approximately independent identically normally distributed, and we present this in Table 8. On the grounds that the nulls of the VAR residual Portmanteau tests for autocorrelations have no residual autocorrelations up to lag *h*, we confirm that the residuals of the VAR model do not exhibit serial correlation at the 5% level, and thus the optimum number of lag periods is 1.

We acknowledge the finding that many macro time series may contain a unit root and the fact that this has spurred the development of the theory of non-stationary time series analysis. If such a stationary linear combination exists, then the non-stationary time series is said to be cointegrated. The stationary linear combination is called the cointegrating equation and may be interpreted as a long-run equilibrium relationship among variables. As explained below, the presence of a cointegrated relation forms the basis of the VEC specification. Consider a VAR of order p. The test hypothesis is formulated as the restriction for the reduced rank of $\Pi = H_0(r) = \Pi = \alpha\beta'$. At the same time, the possibility for substantial technological change and financial liberalization over the past decades is included in the time trend. Thus, we use Johansen's (1988) maximum likelihood approach for our vector autocorrelation model to perform the cointegration tests.

Table 9 reports the results from testing for the number of cointegrating vectors based on the maximum eigenvalue (λ_{max}) and the trace (TRACE) of the stochastic matrix in the multivariate framework that we perform. Table 9 also presents the 5% critical values, which are limited due to the small sample size (Pesaran et al., 2001). Table 9 shows that both tests suggest the existence of one cointegrating vector (r=1) driving the series with common stochastic trends in the data. These results suggest the presence of one cointegrating vector in the four regional house prices, which indicates the existence of a long-run stable relationship for them in Taiwan with quarterly data for the 1993Q1-2009Q2 period. The results of the cointegrated coefficients of the long-run relationship equation are as follows:

$$LTA_t = 0.2685 * LTCY_t + 0.0215 * LTH_t - 0.5990 * LTK_t - 0.0002 * trend \,. \qquad (6)$$

Equation (6) expresses that expansive house prices in LTCY and LTH led to house prices in LTA to increase. What caused this? An important share of housing market movements is related to business cycles. Davis and Heathcote (2006) demonstrate that, in the U.S., residential investment leads the cycle, while non-residential investment lags the cycle. Jud and Winkler (2002) conclude that, applying city level data, real housing price increases are strongly affected by population growth and real changes in income, construction costs, and interest rates. The regional house prices in these two areas, LTCY and LTH, are in northern Taiwan and the suburban areas of Taipei City (the most important economic center in Taiwan), while LTA is in central Taiwan and not far away from the two areas LTCY and LTH. When the business cycles hit the economic center areas of LTCT and LTH, the result is that the regional house prices in these two areas change first and the housing price movements diffuse to the neighboring area of LTA, which is the same as the findings of the above research.

As to expansive house prices in LTK in southern Taiwan, they cause the house price in LTA to decrease. The regions LTA and LTK are centered around traditional manufacturing industries and their similar economic developments also make the two regions' housing be substitutes. In light of the value of elasticity for other regional house prices on LTA, the elasticity of LTK on LTA is the largest, showing that the impact from LTK is much higher than other regions.

What causes the long-run equilibrium of these regional house prices to exist? Ashworth and Parker (1997) find a cointegration between regional house prices, if the regional forcing variables are cointegrated. Meen (1999) indicates that four factors - migration, equity transfer, spatial arbitrage, and spatial patterns in the determinants of house prices - may lead to convergence or cointegration between regional house prices. In the long run the time trend has a low impact on LTA, as indicated by the low value of elasticity for the time trend of -0.002. This shows that the decreasing house price in LTA was one result of Taiwan losing its competitiveness in traditional manufacturing industries after the period of the 1990s.

In order to perform tests on the values of the coefficients for the whole cointegrated vector model, we use Johansen and Juselius' (1990) approach to determine the effect of each variable in the long run during the study period. From the test results shown in Table 10, we have evidence that the effects of LTCY and LTK for equation (6) are significant at the 5% level. Next, it is well known that if the variables are weakly exogenous, then inferences about the parameters of house prices from the conditional distribution are equivalent to inferences about the joint distribution. We adopt the likelihood ratio test of Johansen (1992) to investigate whether or not the variables within the system are characterized by weak exogeneity.

We now examine the existence of weak exogeneity by imposing some linear restrictions on the adjustment coefficients. Since we have identified one cointegrating relationship, we conduct the weak exogeneity test under the assumption of the rank of 1. The test statistics are asymptotically distributed as $\chi^2(1)$ under the null hypothesis of the existence of weak exogeneity. The test results presented in Table 11 reveal that weak exogeneity cannot be rejected for LTCY, LTH, and LTK. Against this, LTA is significantly different from zero, indicating a unidirectional causal relationship running from the other three regional house prices (LTCY, LTH, and LTK) to Taichung (LTA). Thus, Taichung's regional house price is not a fixed exogeneous variable which is set to grow at an exogenous rate.

Why can the other three regional house prices lead the house price of LTA? In light of its geographical position in central Taiwan, LTA is not far away from the other three neighboring areas of LTCY and LTH in northern Taiwan and LTK in southern Taiwan. Through the influence of regional equity transfer or a spatial arbitrage diffusion process, the other three regional house price shocks spread to LTA and cause a unidirectional causal relationship running from the other three regional house prices to LTA. In other words, house price changes in either LTCY, LTH, or LTK lead to housing price movements in LTA. This is an example in which price movements seem to diffuse first from the other three neighboring areas to the geographical center of LTA. The regional house price efficiently diffuses among these regions, which is also caused by the housing market operating on free market principles, because Taiwan's government has a much smaller role in the housing sector. The pre-sale system in Taiwan also can expand the diffusion of regional house prices.

5. CONCLUSIONS and IMPLICATIONS

Most previous research studies on the analysis of stationarity characteristics of regional house prices apply the conventional unit root test, with few studies using panel unit root tests. Recent literature suggests that panel-based unit root tests have higher power than unit root tests based on individual time series. It is well-known that the traditional ADF-type unit root tests suffer from the problem of low power in rejecting the null of non-stationarity of the series, especially with short-span data. The Panel SURADF test developed by Breuer et al. (2001) shows possible cross-sectional effects and identifies how many and which regional house prices within the panel contain a unit root.

The main purpose of this paper is to investigate the stationarity properties and long-run relationship of regional house prices in Taiwan with quarterly data covering 1993Q1–2009Q2. We apply the Panel SURADF test developed by Breuer et al. (2001) to re-examine previous studies of the stationarity of regional house prices. Measuring the half-lives for regional house prices, we provide an analysis of short-run adjustments and the mean reversion process. Conditional upon finding that the variables are I(1), we further use Johansen's (1988) cointegration test to investigate the convergence among regional house prices of I(1). Finally, we employ the weak exogeneity test, as proposed by Johansen (1992), to see whether the variables in the system are characterized by weak exogeneity or not.

From an empirical perspective, the order of integration for the variables has critical implications for the appropriate modeling of data. The permanent versus transitory nature of shocks is related for theoretical models that aim at being consistent with the actual data generating process of the series. Moreover, proper characterizations of the unit root properties of regional house prices are essential in econometric modelling. For instance, in testing for causality between different regional housing activities, a precondition is that both variables need to be integrated of order one (characterised by a random walk). For policy-makers and finance professionals, Diebold and Kilian (2000) also propose that pre-testing for unit roots before applying forecasts yields superior forecasting performance as opposed to the alternatives of working always with differenced series or always using level series. For the above factors, it is worth it to re-examine the stationarity of regional house prices by the Panel SURADF test. According to our knowledge, this is the first fully fledged panel data unit root analysis of regional house prices in Taiwan encompassing this enlarged set of 6 regions. Our main findings are as follows.

First, the results of applying the Panel SURADF test show that six regional house prices in Taiwan are a mixture of I(0) and I(1) processes, and that the traditional panel unit root tests could prompt misleading inferences. The house prices in the Taiwan area and Taipei City have I(0) processes, which reveal that house prices in those respective targets should return to their trend path over time. By contrast, the house price shocks stemming from the other four regions in Taiwan eventually have permanent effects on regional housing markets. This implies that the difference between house prices in these four regions and those elsewhere tend to grow larger and larger over time, with no tendency for the disparity to settle down at an equilibrium level. Thus, the result shows that the stationarity properties of regional house prices are dependent on the respective regions' structure and properties.

For the estimated half-lives of Taipei City, the degrees of mean reversion are greater than those for the overall Taiwan area. The above results thus indicate that applying a more sophisticated testing methodology can reverse findings derived from employing a previous basic approach. A unit root in regional house prices is tantamount to shocks having a permanent effect. This may be an upshot of real factors, such as technology shocks, contributions to economic fluctuations, or aggregate demand and supply shocks having a permanent effect on house price levels, whereas when the impact exhibits temporary (I(0) properties), it might be possible to forecast future movements in the price series based on past behavior. Governments should thus pay attention to long-run trends in the price series and should not adopt excessive targets or interfere with them in the short run. We also consider that once the stationarity can enable government to estimate the relationship between house prices and macroeconomics, this results in higher accuracy and will not yield spurious regressions. Thus, a proper characterization of the unit root properties of regional house prices is essential in econometric modeling. For instance, in testing for Granger causality or estimating the relationship between regional housing prices, a precondition is that both variables need to be integrated of order one. Given that good policy-making typically depends on sound economic forecasts, appropriately modeling the nature of the series is invaluable to forecasters.

Using Johansen's (1988) cointegration test, our findings further provide substantive evidence in favor of the existence of a long-run equilibrium cointegrated relationship among the four regions of Taipei County, Taoyuan-Hsinchu, Taichung, and Tainan-Kaohsiung, indicating that long-run relationships do seem to exist within these four regional house prices in Taiwan. This may ultimately be useful for predicting the long-run tendencies of Taiwan's regional housing market in the presence of macroeconomic shocks. In light of the five regional house prices in Taiwan, there is a long-run equilibrium among all regions except Taipei City, implying a diffusion of regional house prices among each regional house market except Taipei City. Taipei City, the capital and main economic center in Taiwan, has benefitted from the status of a regional global city and the node of a high-technology center (Wang, 2003; Chien, 2010), resulting in much higher housing prices and a different trend path from other regions in Taiwan. As Taipei City is not cointegrated with other regions, this finding has prompted MacDonald and Taylor (1993) to talk of weak segmentation between Taipei City and the other regional house markets.

Finally, the results of the weak exogeneity test indicate the uni-directional causality relationships from the three regions of Taipei County, Taoyuan-Hsinchu, and Tainan-Kaohsiung to Taichung. There are obvious ripple effects of regional house prices that exist from the other three regions to LTA. Moreover, there is no causal relationship between any one of the northern regions (LTCY, LTH) and the southern region (LTK), showing the absence of causality of house prices between the two non-neighboring areas. The empirical results are the same as those in Clapp et al. (1995) and Dolde and Tirtiroglu (1997) – there is spatial diffusion of housing price changes between neighboring areas, but there is none across non-neighboring areas. However, the above results only confirm the presence of the ripple effect, but do not explain it. A more structural model is required further (MacDonald and Taylor, 1993).

Our analysis shows strong evidence overall of regional house price convergence and hence market integration for four regions (Taipei County, Taoyuan-Hsinchu, Taichung, and Tainan-Kaohsiung) in Taiwan's housing markets. Such a finding supports the view that Taiwan's transition to an efficient market economy has been quite successful in respect of its housing markets, which is caused by the housing market operating on free market principles, as Taiwan's government has a much smaller role in the housing sector. The pre-sale system in Taiwan also brings about an efficient diffusion of regional house prices. Furthermore, the choice of where to introduce products and marketing initiatives appears to have relatively little importance.

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References

Akaike, H. 1969. Fitting Autoregressions for Prediction. *Annals of the Institute of Statistical Mathematics* 21: 243-247.

Akaike, H. 1974. A New Look at the Statistical Model Identification, *IEEE Transactions on Automatic Control* AC-19: 716-23.

Alexander, C., and M. Barrow, 1994. Seasonality and Cointegration of Regional House Prices in the UK. *Urban Studies* 31: 1667–1689.

Ashworth, J., and S. Parker, 1997. Modelling Regional House Prices in the UK. *Scottish Journal of Political Economy* 44: 225–246.

Breitung, J. 2000. The Local Power of Some Unit Root Tests for Panel Data, in: B. Baltagi, eds., *Advances in Econometrics, Nonstationary Panels, Panel Cointegration, and Dynamic Panels* 15: 161-178.

Breuer, J.B., R. McNown, and M.S. Wallace, 2001. Misleading Inferences from Panel Unit-Root Tests with an Illustration From Purchasing Power Parity. *Review of International Economics* 9(3)*:* 482–493.

Breuer, J.B., R. McNown, and M.S. Wallace, 2002. Series-Specific Unit Root Tests With Panel Data. *Oxford Bulletin of Economics and Statistics* 64(5): 527–546.

Chang, C.N. and C.W.R. Ward, 1993. Forward Pricing and the Housing Market: the Pre-sales Housing System in Taiwan. *Journal of Property Research* 10: 217-227.

Chen, M.C. 2003. Time-Series Properties and Modelling of House Prices in Taipei Area: An Application of the Structural Time-Series Model. *Journal of Housing Studies* 12(2): 69-90.

Chen, M.C., I.C. Tsai and C.O. Chang, 2007. House Prices and Household Income: Do They Move Apart? Evidence from Taiwan. *Habitat International* 31: 243-256.

Cheung, Y.W. and K.S. Lai, 1993. Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration. *Oxford Bulletin of Economics and Statistics* 55(3): 313–328.

Chien, M.S. 2010. Structural Breaks and the Convergence of Regional House Prices, *Journal of Real Estate Finance and Economics* 40(1): 77-88.

Chiu, R.L. 2002. Testing the Purchasing Power Parity in Panel Data. *International Review of Economics and Finance* 11: 349–362.

Chiu, L.H. 2010. Government Intervention in Housing: Convergence and Divergence of the Asian Dragons. *Urban Policy and Research* 26: 249-269.

Choi, I. 2001. Unit Root Tests for Panel Data. *Journal of International Money and Finance* 20: 249–272.

Clapp, J.M., W. Dolde, and D. Tirtiroglu, 1995. Imperfect Information and Investor

Inferences from Housing Price Dynamics. *Real Estate Economics* 23: 239-269.

Cook, S. 2003. The Convergence of Regional House Prices in the UK. *Urban Studies* 40(11): 2285–2294.

Cook, S. 2005. Regional House Price Behaviour in the UK: Application of a Joint Testing Procedure. *Physica A* 345: 611–621.

Davis, M., and J. Heathcote, 2001. Housing and the Business Cycle. *International Economic Review* 46: 761-784.

Dickey, D. and W. Fuller, 1979. Distribution of the Estimators for Autoregressive Time Series with a Unit Root. *Journal of the American Statistical Association* 74: 427–431.

Diebold, F.X. and L. Kilian, 2000. Unit Root Tests Are Useful For Selecting Forecasting Models. *Journal of Business and Economic Statistics* 18: 265-273.

Dolde, W. and D. Tirtiroglu, 1997. Temporal and Spatial Information Diffusion in Real Estate Price Changes and Variances. *Real Estate Economics* 25: 539-565.

Elliott, G., T.J. Rothenberg, and J.H. Stock, 1996. Efficiency Tests For An Autoregressive Unit Root. *Econometrica* 64: 813-836.

Elliott, G., T.J. Rothenberg, and J.H. Stock, 1996. Efficient Tests for an Autoregressive Unit Root. *Economic Record* 77: 252–269.

Engle, R.F. and C.W.J. Granger, 1987. Cointegration and Error Correction: Representation, Estimation and Testing. *Econometrica* 55: 251-276.

Fisher, R.A. 1933, *Statistical Methods for Research Workers*. Oliver and Boyd, Edinburgh 4th Ed.

Goodman, A. and T. Thibodeau, 1995. Aged-related Heteroskedasticity in Hedonic House Price Equations. *Journal of Housing Research* 6: 25–42.

Hadri, K. 2000. Testing for Stationarity in Heterogeneous Panel Data, *Econometrics Journal* 3(2): 148–161.

Hannan, E.J., and B.G. Quinn, 1979. The Determination of the Order of an Autoregression. *Journal of the Royal Statistical Society* 41: 190-195.

Ho, L.S., Y. Ma, and D.R. Haurin, 2008. Domino Effects within a Housing Market: The Transmission of House Price Changes Across Quality Tiers. *Journal of Real Estate Finance and Economics* 37(4): 299-316.

Holmes, M.J. and A. Grimes, 2008. Is There Long-Run Convergence among Regional House Prices in the UK? *Urban Studies* 45(8): 1531-1544.

Hsueh, L.M. and H.L. Chen, 1998. A Comparison of Household Expenditure by Tenure Choice - Taiwan Evidence. *Journal of Housing Studies* 7: 21-40.

Huang, C.H. 1999. The Housing Price and Quality of Taiwan. *The Proceedings of 1999 Annual Conference of the Chinese Society of Housing Study* 391-400.

Im, K.S., M.H. Pesaran, and Y. Shin, 2003. Testing for Unit Roots in Heterogeneous Panels. *Journal of Econometrics* 115: 53-74.

Johansen, S. 1988. Statistical Analysis of Cointegration Vectors. *Journal of Economic Dynamics and Control* 12(2/3): 231–254.

Johansen, S. 1992. A Determination of the Co-integration Rank in The Presence of a Linear Trend? *Oxford Bulletin of Economics and Statistics* 54: 383–397.

Johansen, S. and K. Juselius, 1990. Maximum Likelihood Estimation and Inference on Cointegration – With Applications to the Demand for Money. *Oxford Bulletin of Economics and Statistics* 52(2): 169-210.

Jud, G.D. and D.T. Winkler, 2002. The Dynamics of Metropolitan Housing Prices. *Journal of Real Estate Research* 23: 29-45.

Kao, C. 1999. Spurious Regression and Residual-Based Tests for Cointegration In Panel Data. *Journal of Econometrics* 90: 1–44.

Kao, C. and M.H. Chiang, 2000. On The Estimation and Inference of a Cointegrated Regression in Panel Data. *Advances of Econometrics* 15: 7–51.

Kim, S.-W. and R. Bhattacharya. 2009. Regional Housing Prices in the USA: An Empirical Investigation of Nonlinearity. *Journal of Real Estate Finance and Economics* 38: 443－460.

Kwiatkowoski, D., P. Phillips, P. Schmidt, and J. Shin, 1992. Testing the Null Hypothesis Of Stationarity Against The Alternative of a Unit Root. *Journal of Econometrics* 54(1-3): 159–178.

Levin, A., C.-F. Lin, and C.S. Chu, 2002. Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties. *Journal of Econometrics* 108: 1–24.

Lin, J.P., and K.L. Liaw, 2000. Labor migrations in Taiwan: characterization and interpretation based on the data of the 1990 census. *Environment and Planning A 32*: 1689-1709.

Ljung, G.M. and G.E.P. Box, 1978. On a Measure of Lack of Fit in Time Series Models. *Biometrika* 65(2): 297–303.

Lűtkepohl, H. 1991. *Introduction to Multiple Time Series Analysis*, New York: Springer-Verlag.

MacDonald, R. and M. Taylor, 1993. Regional House Prices in Britain: Long-Run Relationships and Short-Run Dynamics. *Scottish Journal of Political Economy* 40: 43–55.

Maddala, G.S. and S. Wu, 1999. A Comparative Study Of Unit Root Tests With Panel Data And A New Simple Test. *Oxford Bulletin of Economics and Statistics* 61: 631–652.

Meen, G. 1999. Regional House Prices and the Ripple Effect: A New Interpretation.

Housing Studies 14: 733–753.

Newey, W. and K.D. West, 1994. Automatic Lag Selection in Covariance Matrix Estimation, *Review of Economic Studies* 61: 631-653.

Newey, W.K., and K.D. West, 1987. A Simple Positive Semi-Definite: Heteroskedasticity And Autocorrelation Consistent Covariance Matrix. *Econometrica* 55: 703–708.

Ng, S. and P. Perron, 2001. Lag Length Selection and The Construction Of Unit Root Tests With Good Size And Power. *Econometrica* 69(6): 1519–1554.

Osterwald-Lenum, M. 1992. A Note with Quantiles of The Asymptotic Distribution Of The Maximum Likelihood Cointegration Rank Test Statistics. *Oxford Bulletin of Economics and Statistics* 54(3): 461–472.

Pesaran, M.H., Y. Shin, and R.J. Smith, 2001. Bounds Testing Approaches to the Analysis of Level Relationships, *Journal of Applied Econometrics* 16: 289–326.

Phillips, P.C.B. and P. Perron, 1988. Testing For a Unit Root in Time Series Regression, *Biometrika* 75: 335-346.

Population Census, 2000. The Department of Statistics of Ministry of the Interior for Taiwan, http://www.moi.gov.tw/stat/.

Riddel, M. 2004. Housing-Market Disequilibrium: An Examination of Housing-Market Price and Stock Dynamics 1967-1998. *Journal of Housing Economics* 13: 120-135.

Rossi, B. 2005. Confidence Intervals for Half-Life Deviations from Purchasing Power Parity, *Journal of Business and Economic Statistics* 23(4): 432–442.

Schwarz, G. 1978. Estimating the Dimension of a Model, *Annals of Statistics* 6: 461-464.

Sims, C.A. 1980. Macroeconomics and Reality. *Econometrica* 48(1): 1-48.

Wang, C. 2003. Taipei as a Global City: A Theoretical and Empirical Examination. *Urban Studies* 40: 309 –333.

Yip, N.M. and C.O. Chang, 2003. Housing in Taiwan: State Intervention in a Market Driven Housing System. *The Journal of Comparative Asian Development* 2: 93-113.

Zivot, E. and D.W. Andrews, 1992. Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis, *Journal of Business and Economic Statistics* 10: 251–270.

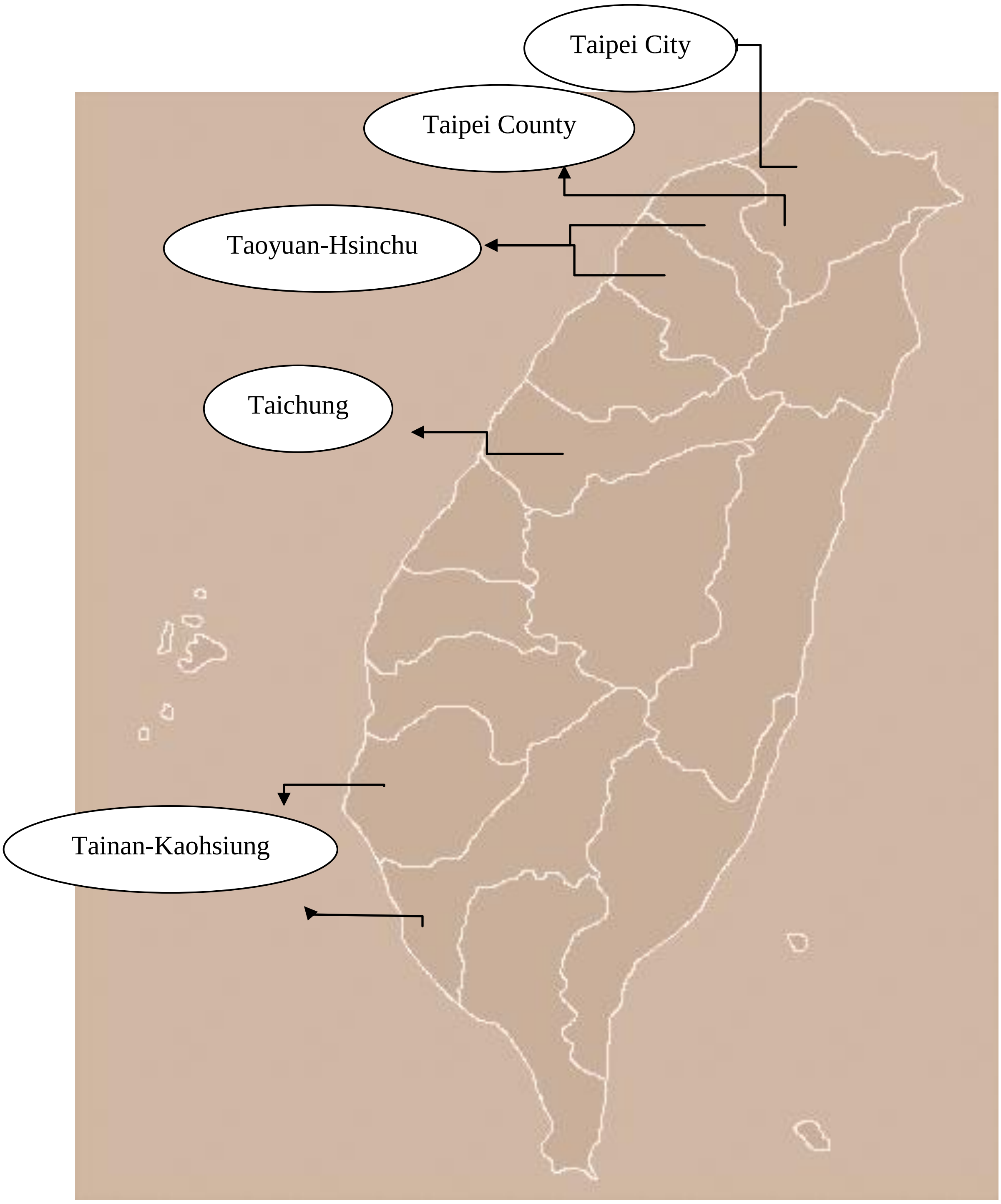

Figure 1. The main metropolitan areas in Taiwan

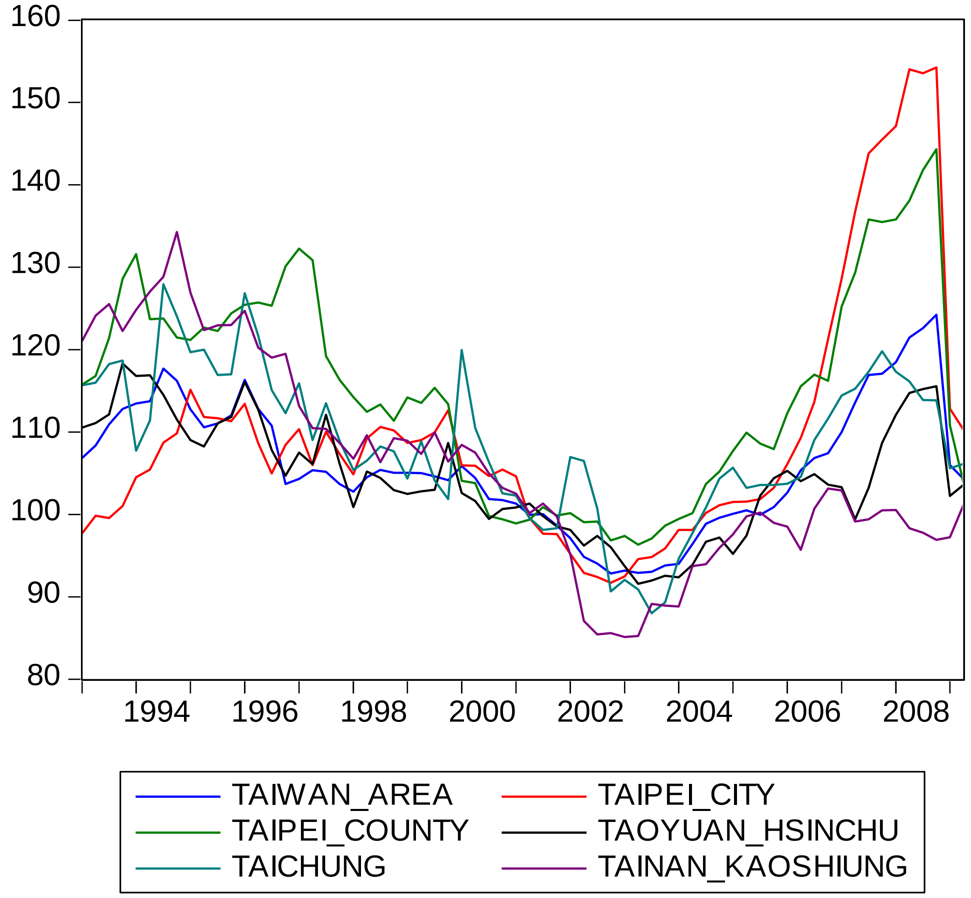

Figure 2. Plots of real regional house prices (1993Q1-2009Q2, 2001 is the base year)

Table 1. Average house price and the ratio of house price to income

Area	Average house price (NT$)	Ratio of house price to income (%)
Taipei City	9,600,000	10.2
Taipei County	5,800,000	6.7
Taoyuan-Hsinchu	5,000,000	6.2
Taichung	5,450,000	6.4
Kaohsiung	4,940,000	5.3
Taiwan Area	6,230,000	7.1

Source: Housing Demand Survey for Fourth Quarter 2008, the Institute for Physical Planning & Information.

Table 2. Univariate unit root tests results of regional house prices

Region	**ADF**	**DF-GLS**	**PP**	**KPSS**	**NP(MZ_{α}^{GLS})**
Taiwan Area	-1.471 (0)	-1.495 (0)	-1.608 (1)	0.203 (6)**	-4.237
Taipei City	-1.593 (0)	-1.625 (0)	-1.904 (3)	0.168 (6)**	-5.055
Taipei County	-1.519 (0)	-1.563 (0)	-1.779 (2)	0.205 (6)**	-4.735
Taoyuan-Hsinchu	-1.963 (0)	-2.005 (0)	-1.992 (3)	0.216 (6)**	-7.375
Taichung	-1.722 (2)	-1.755 (2)	-2.376 (0)	0.192 (6)**	-6.579
Tainan-Kaoshiung	-0.989 (0)	-1.182 (0)	-1.206 (3)	0.190 (6)**	-3.343

Notes: ** indicates significance at the 5% level. DF-GLS and MZ_{α}^{GLS} are unit root tests proposed by Elliot et al. (1996) and Ng and Perron (2001), respectively. The numbers in parentheses are the lag order in the ADF and DF-GLS tests. The lag parameters are selected on the basis of SC. The truncation lags are for the Newey-West correction of the PP and MZ_{α}^{GLS} tests in parentheses. The null hypothesis of the KPSS test examines for I(0), while the null of the remaining four tests examines for I(1).

Table 3. Panel unit root and stationary tests for regional house prices

Method	Statistic
LLC	1.662
Breitung	2.723
IPS	2.244
Fisher-ADF	3.349
Fisher-PP	5.233
Hadri	7.309**

Notes: LLC and IPS represent the panel unit root tests of Levin et al. (2002) and Im et al. (2003), respectively. Fisher-ADF and Fisher-PP represent the Maddala and Wu (1999) Fisher-ADF and Fisher-PP panel unit root tests, respectively. ** indicates statistical significance at the 5% level. Probabilities for Fisher-type tests are computed by using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. The LLC, Breitung, IPS, Fisher-ADF, and Fisher-PP tests examine the null hypothesis of non-stationarity, while Hadri tests the stationary null hypothesis.

Table 4. Panel SURADF tests and critical values for regional house price

Region	SURADF	Critical values		
		0.01	0.05	0.1
Taiwan Area	-3.749 *	-4.516	-3.936	-3.617
Taipei City	-3.587 *	-4.376	-3.842	-3.523
Taipei County	-3.126	-4.239	-3.662	-3.327
Taoyuan-Hsinchu	-2.820	-3.891	-3.314	-2.994
Taichung	-2.185	-3.704	-3.144	-2.867
Tainan-Kaoshiung	-1.567	-3.526	-2.943	-2.657

Notes: * indicates significance at the 10% level. Critical values are calculated using the Monte Carlo simulation with 10,000 draws, tailored to the present sample size. (For details of this simulation, see Breuer et al., 2001.)

Table 5. Panel SURZA tests and critical values for regional house prices

Region	TB	Panel SURADF	Critical values		
			0.01	0.05	0.1
Taiwan Area	1998Q1	-3.729 *	-4.536	-3.935	-3.617
Taipei City	1998Q1	-3.200	-4.343	-3.795	-3.488
Taipei County	1998Q1	-2.835	-4.183	-3.634	-3.326
Taoyuan-Hsinchu	1998Q1	-2.697	-3.916	-3.347	-3.047
Taichung	1998Q1	-2.712	-3.694	-3.141	-2.854
Tainan-Kaoshiung	1996Q2	-1.561	-3.528	-2.944	-2.630

Notes: * indicates respective significance at the 10% level. TB indicates the estimated structural break. Critical values are calculated using the Monte Carlo simulation with 10,000 draws, tailored to the present sample size (for details of this simulation, see Breuer et al., 2001).

Table 6. Estimated half-lives and confidence intervals

Region	β	Half-life (quarter)	Confidence interval at 95%
Taiwan Area	-0.067	9.92	[0, 23.61]
Taipei City	-0.081	8.20	[0, 18.73]

Table 7. VAR lag order selection

Lag intervals	LR	FPE	AIC	SC	HQ
1	289.25 #	$1.45*10^{-12}$ #	-15.92#	-15.22#	-15.64#
2	22.48	$1.60*10^{-12}$	-15.82	-14.56	-15.33
3	22.13	$1.73*10^{-12}$	-15.75	-13.94	-15.04
4	11.38	$2.35*10^{-12}$	-15.48	-13.11	-14.55

Notes: # indicates lag order selected by the criterion.

Table 8. V AR model residuals' portmanteau test for autocorrelations

Model	Number of Lagged Periods	LM(2)	LM(4)	LM(6)
LTA, LTCY, LTH, LTK, trend	1	16.124 [0.444]	9.679 [0.883]	14.393 [0.569]

Note: Figures in [ ] are p-values.

Table 9. Johansen's cointegration test

	λ_{max}		TRACE	
Model	statistics	5% critical value	statistic	5% critical value
$r = 0$	39.479**	34.190	76.823**	67.997
$r = 1$	22.871	27.489	37.342	45.684
$r = 2$	12.483	20.378	14.470	27.541
$r = 3$	1.987	13.325	1.987	13.325

Notes: ** denotes significance at the 5% level, and 5% finite sample critical values are constructed from the asymptotic critical values from Osterwald-Lenum (1992) using the method of Cheung and Lai (1993). r is cointegration rank.

Table 10. Cointegration vector coefficient significance test

	LTCY	LTH	LTK
Statistic	4.460**	0.006	8.788**
P-value	[0.034]	[0.940]	[0.003]

Notes: LR test statistic is obtained by means of the $\chi^2(r)$ test; the figures inside [ ] are the p-values. ** denotes significance at the 5% level.

Table 11. Weak exogeneity test

	LTA	LTCY	LTH	LTK
Statistic	15.511**	0.972	0.006	2.072
P-value	[0.000]	[0.324]	[0.980]	[0.150]

Notes: LR test statistic adopted is the $\chi^2(r)$ test statistic. The figures in [ ] are p-values. ** denotes significance at the 5% level.