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Cointegration of Matched Home Purchases and Rental Price Indexes – Evidence from Singapore

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Abstract

This paper exploits the homogeneity feature of the Singapore private residential condominium market and constructs matched home purchase price and rental price series using the repeated sales method. These matched series allow us to conduct time series analysis to examine the long-term present value relationship in the housing market. Three key findings are obtained. First, we fail to establish a cointegrating relationship between the home purchase price and rental price based on nationally estimated indexes. Second, area-specific indexes demonstrate strong cross-correlations, invalidating the use of first generation panel unit root tests that ignore these cross-correlations. Third, Pesaran's CIPS test indicates that the unit root hypothesis is rejected for the first difference of both indexes. We also do not reject the hypothesis that area-specific home purchases and rental price indexes are cointegrated with a cointegrating vector (1,-1).

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Keywords: Cointegration; Housing Market; Purchase and Rental Price; Market Efficiency

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1. Introduction

It is well established that housing markets exhibited exaggerated cyclical patterns, which is especially highlighted by the 2007 U.S. financial and housing crisis. As a first step in understanding housing price dynamics, several measures of fundamentals have been proposed with which to compare house prices. One of the most widely adopted measures is the present value of rents.[1] However, a long-standing issue with using the deviation of price to rent as a proxy for mispricing resides in the appreciable difference in the *quality* of units that are transacted on the housing sales market and the housing rental market respectively. This paper exploits the homogeneity feature of the Singapore private residential condominium market to match the quality of housing. We construct home purchase price and rental price series for nearly identical units on this market based on the repeated sales method. This approach provides an opportunity to explore and better understand various implications of the housing price present value relationship both in the short run and in the long run.

There has been an extensive literature examining the cost of owning a home relative to renting using the present value model. This model is also referred to as the user cost model, which defines the equilibrium relationship between housing rents and prices, after taking into account favorable tax treatments given to owner occupied properties and mortgage interest payments. Beginning with Poterba (1984), many authors have compared the user cost of owner-occupying with the cost of renting to assess potential mispricing and generate implications on the efficiency of the housing market (Meese and Wallace, 1994; Clark, 1995; Chen, 1996; Leamer, 2002; Crone et al. , 2004; Krainer and Wei, 2004; Cutts et al., 2005;

[1] Other measures used to gauge housing prices include comparing house prices to the underlying construction cost or corresponding economic fundamentals, such as income, population, etc. (See Poterba, 1991; Rosenthal, 1999; Case and Shiller, 2003; McCarthy and Peach, 2004; Gallin, 2006; Holly et al., 2010, to mention a few.)

Himmelberg et al., 2005; Ayuso and Restoy, 2006; Davis et al. , 2008; Gallin, 2008; Campbell et al., 2009; Mayer and Sinai, 2009; Ambrose, Eichholtz, and Lindenthal, 2013; Feng and Wu, 2015, etc). This pricing strategy is similar to the dividend discount model for the stock market, except that the yield to housing is the rent-price ratio. Campbell et al. (2009), in particular, apply the dynamic Gordon growth model which decomposes the rent-price ratio into the expected present discounted values of rent growth, real interest rates, and a housing premium and find similar housing dynamics to those found for stocks and bonds.

The present value model has its strength in providing a convenient framework to consider the impact of the user cost on house prices as well as to explore potential mispricing in the housing market. However, researchers often use different price and rent indexes under the assumption that the rent index is a good proxy for the rent that might be paid for an equivalent owner-occupied property. For example, in the Meese and Wallace (1994) study, the characteristics of the rental sample do not exactly match that of the owner-occupied sample. Gallin (2008), as another example, uses the housing price index published by the Federal Housing Finance Agency, which is based on a sample of housing units unmatched to those included in the rental shelter index from the Consumer Price Index. As Glaeser and Gyourko (2007) point out, such comparison is inaccurate given that dwellings included in the price indexes do not match the dwellings in the rental indexes. In fact, the owner-occupied houses are often better maintained than rental houses (Shiling, Sirmans, and Dombrow, 1991). With comparatively poorly maintained rental units, the time series path of the rental price indexes may vary significantly from the implicit rents of the owner-occupied units.

The lack of homogenous units transacted on both the property sales market and the rental market has been recognized and addressed in the literature in various ways. For instance, Smith and Smith (2006) and Hill and Syed (2012) make use of owner-occupied houses that have comparable characteristics to those that are also available for rent. Garner and Verbrugge (2009) rely on the U.S. Consumer Expenditure Survey to compare self-reported rents and house values of the same house. A more recent study is Bracke (2014) which isolates properties that were both sold and rented out within six months and measure their rent-price ratios. However, due to data availability, these studies often focus on exploring matched data at the micro-level to shed light on cross-sectional variation in rent-price ratios.[2] They lack the data capacity to construct both rental and price series for an extended and continuous period of time. The ability to construct time series data using matched rental and owner-occupied units is, however, important for a better understanding of the long-term cointegrating relationship between prices and rents and their short-term price dynamics.

This paper extends the literature by constructing matched home purchases and rental price quarterly index series (from 2000Q1 to 2014Q4) using transaction-level data from the Singapore private condominium market. All units within each residential condominium project are homogenous given that they are all fully furnished units with the same interior design, the same type of furnishing, the same major electrics, and the same outdoor facilities. This means that, after adjusting for observed characteristics, we have essentially identical units that are transacted on both the property sales market and the corresponding rental market at the same time. This feature enables us to construct both the purchase price index sequence and the rental price index sequence for a sample of residential housing projects that have their units both purchased by a homebuyer and rented out to a tenant at a certain point

2 Although Hill and Syed (2012) managed to construct 9 years of quality-adjusted price-rent ratios, they are of relatively low frequency, which is not sufficient for rigorous time series analysis for long-run relationships.

in time. More importantly, the richness of the data allows us to construct both quarterly purchase price and rental time series indexes based on matched units for twelve years of time span and for separate geographic areas. This provides valuable opportunities for rigorous time series analysis, taking into consideration potential cross-sectional heterogeneity and spatial correlations.

The construction of the time series indexes makes use of the repeated sales method as proposed in Bailey, Muth, and Nourse (1963) and later generalized and popularized by Case and Shiller (1989). The idea is to rely on a set of units (or residential projects in this case) that have been transacted (or rented out) more than once during the sample period. The percentage change in house prices (or rental prices) between two sale dates is regressed on a set of dummies associated with the quarter of the turnover. Attributes of the home and their shadow prices are assumed to be unchanged between turnover dates and, therefore, drop out of the model. This method allows us to construct time series variables not only for *national* home purchase-rental price indexes (based on the national sample), but also for *area-specific* home purchase-rental price indexes (based on the area-specific sample). The latter will be particularly useful in addressing cross-section heterogeneity and spatial dependence in the housing market, as will become apparent later.

We first use the constructed home purchase and rental price indexes at the national level to test for the long-term cointegrating relationship of the log real home purchase and rental prices, as suggested by the theory. Unit root tests have been employed to test the stationarity of both series. We find that at the national level, both indexes generally follow the I(1) process. However, the cointegration test fails to reject the null of a unit root in the residuals of the regression of real house purchase prices on real rental prices. This is likely caused by the short time span of the data that we consider in this paper or potential cross-sectional heterogeneity that may bias the results. It might also be caused by the possibility that the weighted national average of purchase and rental price series disguises the underlying cointegrating relationship that might be otherwise present at a more refined geographic level.

To cope with the above-mentioned problem, we construct separate home purchase and rental price index pairs for ten heterogeneous areas in Singapore based on a sample of repeatedly transacted residential projects in each of these areas. This allows us to utilize the panel structure of the data and to take into consideration possible cross-sectional dependence of the time series across these heterogeneous areas. To elaborate on the latter, for example, a set of common shocks to the embedded user cost of owner-occupying may affect each area simultaneously and contribute to cross-sectional correlation of both the purchase and the rental price indexes across these areas. We apply the common correlated effects (CCE) estimator of Pesaran (2006) which allows for unobserved common factors to be possibly correlated with area-specific regressors. This estimator is consistent under heterogeneity and cross-sectional dependence.[3]

Three key findings are obtained for area-specific home purchase and rental prices. First, the cross-sectional dependence (CD) test statistics of (Pesaran, 2004) show that the cross-correlations are statistically significant, and thus invalidate the use of first generation panel unit root tests, such as the Im, Pesaran and Shin (2003) IPS test, which does not allow for cross-sectional dependence (see Breitung and Pesaran, 2008). Second, allowing for second-generation panel unit root tests that take into consideration cross-section dependence, like the

[3] The CCE procedure also copes with the presence of spatial effects (Pesaran and Tosetti, 2010). This is because spatial dependence is dominated by the common factor error structure that underlies the CCE estimator.

Pesaran's CIPS test, we find that the unit root hypothesis is rejected for the first difference of the log real house purchase and rental price indexes, respectively (Pesaran, 2007). This result is robust to the choice of the lag order underlying the cross-sectionally dependent augmented Dickey-Fuller (CADF) regressions. This result is also invariant to whether trends are taken into account. Third, our panel cointegration test suggests that, when a time trend is included, area-specific home purchases and rental prices are cointegrated with a cointegrating vector of (1,-1). This suggests that in the long run, home purchase prices do not significantly deviate from the corresponding rentals and any persistence in present value errors is transitory. This is consistent with the long-run implications of housing market efficiency.

The reminder of the paper is organized as follows: Section 2 introduces the institutional background of the Singapore private condominium market. Section 3 provides a theoretical review of the present value model. Section 4 discusses the data and the construction of home purchases and rental price indexes. Section 5 presents the empirical findings. Section 6 concludes.

2. Singapore Private Condominium Market

In general, residential properties in Singapore can be grouped into three major categories: private apartments/condominiums, landed property, and public housing locally known as Housing and Development Board (HDB) flats.[4] Among all three categories, public housing is the most popular dwelling in Singapore. Based on the 2005 General Household Survey, about 82% resident households live in HDB dwelling. Condominium and apartment flats are occupied by 12% of resident households and the rest live in landed properties.[5] Within the

[4] HDB is the statutory board of the Ministry of National Development.
[5] http://www.singstat.gov.sg/statistics/browse_by_theme/ghs.html.

private property residential market, condominium housing is the largest. It accounts for 38% of the total available private residential stock. Together with apartments, non-landed properties constitute two thirds of the accumulated stock. 75% of the condos are owned by Singaporeans. 56% of these units are used for personal living, while the rest are rented out.[6]

Compared to the other two segments of the housing market in Singapore, private condo units are much more homogenous within each residential project. Typically, private condo projects have their own security guard and enclosed car parks and all condo units are fully furnished. Each unit within the same project has the same type of furnishing (window design, wall painting, floor lamination, built-in closets, built-in kitchen cabinet, etc.). They also come with the same basic household appliances (air-conditioner, washing machine, microwave, kitchen oven and hood, refrigerator, etc.). They are often equipped with various facilities such as swimming pool(s), Jacuzzi, tennis court, gym, squash court, basketball court, children playground, clubhouse, BBQ area, etc. Although the design and the type of amenities vary from one project to another, all households within the same residential condominium project have access to all its amenities.

Given the full extent of furnishing, homebuyers typically do not engage in large furnishing and decorative schemes before move-in. This feature makes the rental units and the owner-occupied units much more comparable in terms of their qualities. It is this homogeneity feature of the Singapore private condominium market that allows us to find almost identical units that are transacted on both the sales market and the rental market. Accordingly, the home purchase price index and rental price index can be constructed for the

[6] Comparable market size of both the owner occupied and the renter occupied in the private condominium market is considered as another nice feature that facilitates our analysis of both the purchase price and rental price at the same time.

same sample of residential condominium projects over time, which allows for the analysis of the long-term cointegrating relationship between real purchase prices and rents.

3. Present Value Model

One of the implications of an efficient housing market is the present value model which relates the real home purchase prices to the corresponding real rental prices. In an efficient market, the real expected rate of return from owner occupation needs to be set against the homeowner cost of capital k_t. Denote the real house purchase price at the beginning of period t by P_t, and the real rental cost of the same house over period t by R_t. The real rate of return from owning the house over period t is given by: $(P_{t+1} + R_{t+1} - P_t)/P_t$. This expression abstracts from housing service consumption, transaction costs, depreciation, and other costs of home ownership. [7]

The one period arbitrage condition for the asset market equilibrium is given by $\frac{\mathrm{E}(P_{t+1} + R_{t+1} | \mathcal{F}_t)}{P_t} - 1 = k_t$, or equivalently, $P_t = \mathrm{E}(P_{t+1} + R_{t+1} | \mathcal{F}_t)/(1 + k_t)$, where $\mathcal{F}_t$ is the information set available at time t. Repeated use of the above equation leads to bubble-free real house prices as determined by the sum of a discounted stream of future rental flows. The solution simplifies considerably under $k_t = k$:

$$P_t = \sum_{j=1}^{\infty} \left(\frac{1}{1+k}\right)^j E\left(R_{t+j} \middle| \mathcal{F}_t\right),$$

[7] We follow similar framework as used in Holly et al. (2010) and assume that the user cost of housing capital can be generalized and represented by a time-varying interest rate. The long-run relationship between housing rents and purchase prices holds as long as the generalized interest rate is sufficiently large relative to the rental growth rate. Housing asset in this case is considered as a pure investment vehicle. To take into consideration the consumption needs of home ownership, the net benefit should be described as $P_{t+1} - P_t + S_t$, where S_t is the real value of housing services. To arrive at the same cointegration condition, we shall assume that $S_t = \alpha_t R_t$, where $\alpha_t > 1$ and is a stationary process.

which is equivalent to

$$\frac{P_t}{R_t} = \sum_{j=1}^{\infty} E\left(\prod_{s=1}^{j}\left(\frac{1+g_{t+s}}{1+k}\right)\middle|\mathcal{F}_t\right),$$

letting $g_t = \Delta\ln(R_t)$. Therefore, under the assumption that $g_t = g + \varepsilon_{gt}$, $\varepsilon_t \sim \mathrm{i.i.d.}(0, \sigma_g^2)$ and that k is large relative to the growth rate of house rental prices, the price-rent ratio, P_t/R_t, will be a stationary process. In particular, $p_t = \ln(P_t)$ will be cointegrated with $r_t = \ln(R_t)$ with the cointegrating vector (1, -1).

In the empirical work that follows we allow for institutional changes in the real homeowner cost of capital variable(k_t). This is because a decline in nominal interest rate, for example, may reduce the homeowner cost of capital. Thus higher housing prices can be associated with the same flow of rents. This can be addressed by including a set of time-varying common factors while applying CCE estimation in a panel data setting.[8] The estimated coefficients obtained following the CCE procedure are consistent given a fixed number of time-specific factors that may or may not be stationary (Kapetanios et al. 2009).

4. Construction of Price Indexes

4.1 Data

The empirical analysis to follow relies on two primary data sources. The first data set is the Real Estate Information System (REALIS) maintained by the Urban Redevelopment Authority of Singapore (URA).[9] The REALIS database provides timely and comprehensive information on Singapore private property market, including private residential, commercial and industrial sectors. The database can be accessed through subscription services. We

[8] Note that the inclusion of cross sectional averages, a feature of CCE, also proxies for other short-term time-specific unobserved influences.
[9] https://spring.ura.gov.sg/lad/ore/login/index.cfm

focus on the private residential condominium market alone to make use of its unique homogeneity feature. For this specific market sector, we obtain information on the floor-area-adjusted median purchase price for each condo project in each quarter from 2000 quarter 1 to 2014 quarter 4. To calculate the floor-area-adjusted price, we first divide the transaction price by its corresponding floor area. We then take the median of the floor-area-adjusted price among all the transactions within a quarter and for a particular condominium project. The median floor-area-adjusted prices are, in this sense, specific to each condo and each quarter. We only keep records of projects that have at least three transactions during a quarter to reduce noise. This is also to better match how the rental data are reported as shown below.

The second dataset contains the corresponding rental information which is also provided by the URA but through a public portal. On the URA's website, information on floor-area-adjusted median rental price of each private condo project can be obtained for all quarters starting from the year 2000 quarter 1.[10] The floor-area adjustment is done in the same way as for the purchase price. The median rental is also obtained as the median floor-area adjusted rental price specific to each condo in each quarter. The slight difference is that it is reported by URA only for projects within which at least ten rental contracts have been signed during the specific time period to ensure the representativeness of the statistics. Accordingly, as mentioned above, we modify the sales transaction data in a similar way so as to be consistent with the way in which rental data are reported.

To ensure that the purchase price and the rental price pertain to units that share almost identical characteristics, we further restrict our sample to those projects that have both valid

[10] https://www.ura.gov.sg/realEstateIIWeb/resiRental/search.action

median rental and valid median purchase price recorded at the same time. In this way, we only focus on projects whose units have been transacted relatively more frequently on both the rental market and the purchase market. To construct time series data on both price indexes, we rely on the repeated sales method to estimate the national home purchase-rental price index pair as well as area-specific home purchase-rental price index pair for a sample of projects that have their floor-area-adjusted median rental/sales prices recorded at least twice on the respective market.

4.2 Repeated Sales Method

We illustrate the construction of home purchase price index based on repeated sales as follows. The rental price index can be constructed in a similar fashion.

As in Bailey, Muth, and Nourse (1963), Case and Shiller (1989), and Case and Quigley (1991), suppose that the price of a home $P_{i,t}$, is observed upon purchase and sale in periods t and $t+\tau$ separately, where $P_{i,t} = e^{\gamma_t} f(X_{i,t}; \beta_t)$ and $P_{i,(t+\tau)} = e^{\gamma_{t+\tau}} f(X_{i,(t+\tau)}; \beta_{t+\tau})$. In these price equations, f is an unknown and non-linear function of the period and unit specific characteristics (X) of home i, and the corresponding shadow price (β) of these characteristics. The elements of X include both structural attributes and characteristics of the neighborhood specific to the house. The terms γ_t and $\gamma_{t+\tau}$ represent the influence of period-specific market conditions that are common to all properties in the geographic market from which the sample of homes are drawn. These terms measure the quality-adjusted price of housing in period t and $t+\tau$.

Suppose now that both X and β are unchanged between sales.[11] We have $P_{i,(t+\tau)} = e^{\gamma_{t+\tau}-\gamma_t} P_{i,t}$. Taking logs and rearranging, we obtain $\log \frac{P_{i,(t+\tau)}}{P_{i,t}} = \gamma_{t+\tau} - \gamma_t$. Taking this equation to the data, we get $\log \frac{P_{i,(t+\tau)}}{P_{i,t}} = \sum_{s=1}^{n} \gamma_s D_{i,s} + \varepsilon_i$, where ε_i is a random error term and $D_{i,s}$ equals -1, 1, or 0 depending on whether the corresponding time index, s, refers to t, $t + \tau$, or other periods.

Following the above-mentioned procedure, we construct home purchase price series and rental price series based on quarterly repeated median rent/price of condominiums. Note that this approach is different from the standard repeated sales procedure, which is based on repeated unit transactions. The adoption of this approach is, first of all, due to data availability: housing in Singapore is relatively newly built and there are a quite limited number of repeated purchase transactions. Although Deng et al (2012, 2014) have proposed a matching procedure to overcome this difficulty by making the best use of information available for purchase transactions, the rental records unfortunately are not available at such detailed transaction level. The second reason that we match pairs at the aggregate level is that we believe it is a valid approach given the homogeneity feature of Singapore condominium markets. In fact, it has been shown by Jiang, Phillips, and Yu (2015) that this method is less susceptible to specification errors than standard hedonic methods and is not

[11] This assumption, although it may not hold, has been imposed quite often in the literature when conducting similar estimation. We argue that this assumption is more likely to be true in our context. This is because we are tracking the median transaction prices at the project level over time, and hence we are more likely to find transaction pairs that are relatively close in time. In fact, our data show that the average time lag in-between a transaction pair is 2.3 quarters. This is much more frequent than the standard repeated transactions at the unit level which often take years to repeat. Given the small time lag in-between transaction pairs, we think it is reasonable to assume that both the housing characteristics and the corresponding shadow prices are unchanged between turnovers.

subject to the sample selection bias that repeated sales index is often criticized for. The above-mentioned paper also finds that the out-of-sample predictions based on this new method are better than both the Case-Shiller index and the hedonic index.

Using the above-mentioned method, we construct matched rental and price indexes both at the national level based on the nation-wide sample and the regional level based on the area-specific sample of repeated housing transactions. The former is plotted in Figure 1, with comparison to the private residential property purchase index and rental index published by CEIC.[12] With both series normalized as 100 in 2000 quarter 1, we see that the rental price series are generally above the purchase price series, which is similarly captured by both self-generated indexes and CEIC indexes. One discrepancy is that the self-generated purchase and rental price indexes seem to have more ups and downs as compared to the corresponding CEIC indexes. The reason could be that we are picking up a sample of projects that are more frequently transacted on the market in producing the self-generated repeated sales indexes. These more frequently traded projects are more likely to be subject to speculative behavior and hence are more sensitive to fluctuations in market conditions.

To better control for potential cross-sectional heterogeneity, we next construct separate home purchase and rental price index pairs for ten heterogeneous areas in Singapore based on a sample of repeatedly transacted residential projects in each of these areas. These are generally small clusters of urban planning areas as defined by the URA.[13] There are 55 urban planning areas in Singapore, spanning five different regions nationwide. Each planning area has a population of about 150,000 and is served by a town center and several

[12] https://www.ceicdata.com/
[13] http://www.ura.gov.sg/uramaps/?config=config_preopen.xml&preopen=Planning Boundaries&pbIndex=1

neighborhood commercial/shopping centers. For the purpose of this study, we group urban planning areas into ten different area clusters (also referred generally as "area" in this context) based on how integrated the neighboring areas are. The clustering of urban planning areas helps to expand the repeated sales sample size for each area to obtain more accurate index measures at all points in time. There areas include the following: The Central Core Area (CC), the Central West Bound Area (CWB), the Central East Bound Area (CEB), the Central West Periphery Area (CWP), the Central East Periphery Area (CEP), the West South Bound Area (WSB), the West North Bound Area (WNB), the North and Northeast Region (NNE), the East South Bound Area (ESB), and the East North Bound Area (ENB). These ten different planning area clusters are shown in Figure 2. The corresponding logarithm of housing purchase and rental price series for each area are plotted in Figure 3, with the value in 2000 quarter 1 normalized as 1.

There is considerable heterogeneity in the time series patterns of both indexes across different areas. For instance, for the central core area, purchase prices are generally above the rental prices especially after 2003, whereas it is less likely to be so for other periphery areas. For the East South Bound (ESB) area, both indexes generally track each other; while for the nearby East North Bound (ENB) area, there is a larger deviation between purchase prices and rental prices. This highlights the importance of controlling for heterogeneity across different areas.

5. Time Series Analysis

5.1 Nation-Wide Series

We begin our analysis by exploring the cointegration relationship of the nationally estimated home purchase and rental price series. Standard time series procedures have been conducted to first examine the stationarity of both indexes. Results are reported in Table 1. As shown in the table, both the log of the real purchase price index and the log of real rental price index fail the unit root test in almost all cases. For the first differences of both series, the unit root hypothesis is rejected in most cases. Evidence suggests that both the log real purchase price index and the log real rental price index likely follow I(1) processes. This leads to the next step of examining the long-run cointegrating relationship of both series.

To test for possible cointegration between the nationally estimated home purchase and rental price indexes, we first obtain the residuals after regressing the purchase prices on rental prices and then test whether the residual series is stationary or not. Results are reported in Table 2, where the existence of unit root in the residual series cannot be rejected for all cases. Hence, we fail to establish the cointegrating relationship between the home purchase price and rental price based on nationally estimated indexes.

The failure of establishing the nation-wide cointegrating relationship between matched home purchase and rental prices is likely caused by the short time span of the data. It may also be caused by the fact that the weighted average price series at the aggregate level disguises the underlying cointegrating relationship that might be otherwise present at the more refined geographic level. To address these concerns, we construct separate home purchase and rental price index pairs for ten heterogeneous areas in Singapore (as defined earlier) and examine the time series properties of these index sequences in a panel structure setting.

5.2 Area-Specific Series

To examine the likely existence of long-term cointegrating relationship between matched home purchase and rental prices in a panel setting, we start with a panel unit root test. One of the most commonly adopted panel unit root tests is proposed in Im et al. (2003), also referred to as the IPS test. This test, however, is not valid when time series are cross-sectionally dependent. To ensure the proper use of the panel unit root test, we first conduct the CD test to check possible cross-correlations in both the purchase price series and rental price series. As expected, the CD test statistics, as reported in Table 3, show that the cross-correlations are statistically significant, which invalidates the use of the IPS test that does not allow for error cross-sectional dependence. Therefore, in what follows we focus on the CIPS test proposed by Pesaran (2007), which follows the CCE procedure and filters out the cross-sectional dependence by augmenting the ADF regressions with cross-section averages.

As reported in Table 4, Pesaran's CIPS panel unit root test convincingly rejects (at the 1% level) the existence of unit root in the first difference of both purchase price and rental price series for all lag orders. The unit root hypothesis, however, cannot be rejected for the levels of both series, with the only exception for the rental index with lag order 1 and with both an intercept and a linear trend included in the underlying CADF regression. The evidence suggests that both purchase price and rental price series likely follow I(1) processes.

To test for the cointegrating relationship between area-specific home purchase and rental price indexes, we estimate a model following the standard setup that allow for heterogeneous slope coefficients, as shown below.

$$P_{it} = \alpha_i + \beta_i R_{it} + u_{it}, \text{where } u_{it} = \sum_{l=1}^{m} \gamma_{il} f_{lt} + \varepsilon_{it}, i = 1,2,\cdots,N; t = 1,2,\cdots T. \quad (1)$$

We assume in the model that u_{it} is captured by a fixed number of factors, f, which may or may not be stationary and a stationary error term, ε (Kapetanios et al. 2009). These factors represent, for instance, time-varying real cost of borrowing, time-specific policy changes, or other macroeconomic factors that are time-specific. To obtain a consistent estimate of β (the cross-sectional average estimate of β_i), we adopt the CCE estimation technique (Pesaran, 2006) which is based on the following cross-section augmented regression

$$P_{it} = \alpha_i + \beta_i R_{it} + d_{i0}\bar{P}_t + d_{i1}\bar{R}_t + e_{it}, \quad (2)$$

where $\bar{P}_t$ and $\bar{R}_t$ denote the cross-section averages of P_{it} and R_{it}. The results are reported in Table 5. The first column of Table 5 reports the basic mean group (MG) estimates which do not take into consideration cross-sectional dependence. The second and the third columns report the common correlated effects mean group (CCEMG) estimates and the common correlated effects pooled (CCEP) estimates, where the latter differs from the former in assuming away heterogeneous slope coefficients. Results indicate that CCE estimates greatly reduce the average error-correlation coefficient, from 0.564 to -0.043(CCEMG) or -0.070 (CCEP). But at the same time, we do not find significantly different coefficients associated with the rental index. The hypothesis that $\beta = 1$ cannot be rejected in all cases. Hence, we proceed by assuming that the long-run cointegration to be tested for is as follows:

$$\hat{u}_{it} = P_{it} - R_{it} - \hat{\alpha}_i \quad (3)$$

where $\hat{\alpha}_i = \frac{1}{T}\sum_{t=1}^{T}(P_{it} - R_{it})$.

The residual defined above can be used to test the null of non-cointegration between the area-specific rental indexes and purchase price indexes. Note that the inclusion of common factors in Eqn. (1) requires that the panel unit root tests applied to $\hat{u}_{it}$ should allow for cross-sectional dependence. Hence, we compute the CIPS panel unit root test statistics for different augmentation and lag orders. Results are reported in Table 6. For the case when both an intercept and a linear trend are included in the underlying CADF regression, we significantly reject (at the 1% level) the null of unit root in the residuals for all augmentation orders. The existence of unit root in the residuals can only be significantly rejected at 10% level with augmentation order 1 in the case when only an intercept is included. Evidence suggests that, especially after taken into consideration the linear trend, the area-specific home purchase and rental prices are cointegrated with cointegrating vector (1, -1).

The national-level violation of – but area-level consistency with – the long-run cointegration between prices and rents can be rationalized by the following considerations. From a statistical perspective, this is likely caused by the potential cross-sectional heterogeneity which cannot be directly controlled for in national-level time series analysis and ends up biasing the results. It could also be that the weighted national average of purchase and rental price series disguises the underlying cointegrating relationship that might be otherwise present at a more refined geographic level. From an economics perspective, these findings are consistent with various discussions on housing segmentation (Goodman and Thibodeau, 1998; Leishman, 2001; Islam and Asami, 2009; etc). In particular, Piazzesi, Schneider, and Stroebel (2015) document that market and search activity varies significantly within narrow geographic areas. In this sense to better account for these cross-sectional variations, housing markets should be examined at a more refined geographic level to uncover any short-term or long-term price dynamics.

The established cointegration and cointegrating vector coefficients help us understand the underlying long-run relationship between the matched home purchase and rental price series. It suggests that in the long run, house prices and rents do not deviate significantly from each other. In other words, price deviations from the present value of rents are mean reverting.

The evidence is not necessarily in conflict with findings of housing market inefficiency in the short run. [14] This is because in the short term, any deviations between the imputed rent and the homeowner cost of capital have to exceed transaction costs before any adjustment in the consumption of housing services would take place. In the long run factors like transaction costs become infinitely averaged out and long-term arbitrage opportunities eventually eliminate any long-term predictability in returns, as shown in Meese and Wallace (1994) and Clark (1995).

Having established the long-term cointegration relationship, we estimate the panel error correction model to uncover the dynamics of the short-run adjustment of real house prices to rental prices. The error correction model is specified as follows:

$$\Delta P_{it} = \alpha_i + \mu_i(P_{it} - R_{it}) + \delta_{i1}\Delta R_{it} + \delta_{i2}\Delta P_{i(t-1)} + \upsilon_{it}, \tag{4}$$

We report CCEMG and CCEP estimates of the parameters, as well as the mean group estimators which do not take into account cross-sectional dependence. Results are summarized in Table 7. The mean group estimates report an error correction coefficient of -0.1102 (0.023), which corresponds to a half-life shock of approximately 5.937. The average cross-correlation of the residuals is 0.315 and the corresponding CD test statistic is 16.071. Both indicate that the MG estimators are likely to be biased due to cross sectional dependence.

The CCEMG and CCEP estimators are reported in column 2 and column 3 of Table 7. The corresponding error correction coefficients are -0.1655 (0.025) and -0.1581 (0.039),

[14] For instance, Case and Shiller (1989, 1990) find evidence of predictability in excess returns. Poterba (1991) argues that none of the conventional asset pricing explanations – changes in the user costs, changes in construction costs, and changes in demographic factors – fully represents house price movements. Both findings are consistent with the implications of housing market inefficiency in the short run.

respectively. This suggests that, after taking into consideration cross-sectional dependence, the real purchase price is more responsive to a rental shock, especially for the case assuming heterogeneous slope coefficients. The half-life corresponding to both error correction coefficients is 3.831 (CCEMG) and 4.028 (CCEP), respectively. The residual cross-sectional dependence is reduced from 0.315 for the MG estimates to -0.097 and -0.106 for the CCEMG and CCEP estimates, respectively.

Results from the error correction model are useful for uncovering short-term housing price dynamics. We find evidence of a half-life that is roughly 3.9 quarters: the time needed to eliminate 50% of the deviation is about one year. This is longer than what Meese and Wallace (1994) have documented for the U.S. In fact, they find that present value errors exhibit persistence with a half-life of about 4 months over the period 1970-1988. This is, however, shorter than Holly et al (2010)'s findings of average half-life for the deviations of price to per capita income for the U.S. being around 3.5 years. Findings of the half-life for Singapore in this paper significantly contrast with those of a recent paper by Ambrose et al. (2013) which shows that the market correction of potential mispricing in Amsterdam may take decades.

6. Conclusion

This paper examines the long-run implications of the present value relationship by drawing on the homogeneity feature of the Singapore private condominium market. This particular market segment consists of residential projects with almost identical units that are available both for sale and for rent. These homogeneous units allow us to construct matched home purchase and rental price index series from 2000 quarter 1 to 2014 quarter 4 for both the nation as a whole and for each specific area.

We conduct both national level time series analysis and area-specific analysis in a panel setting. We obtain the following results. First, we find that at the national level, both indexes generally follow the I(1) process. But the cointegration test fails to reject the null of a unit root in the residuals of the regression of real house purchase prices on real rental prices. In other words, we fail to establish the cointegrating relationship between the home purchase price and rental price based on nationally estimated indexes. Second, area-specific home purchase and rental indexes show significant cross-sectional dependence, which invalidates the use of first generation panel unit root tests, such as the IPS test. Unit root tests based on the CIPS test suggest that the area-specific home purchase and rental prices follow the I(1) process and are cointegrated with cointegrating vector (1, -1).

We think the failure of establishing the long-run cointegration between nationally estimated home purchase and rental prices is likely caused by the short time span of the data or the possibility that the weighted average price series at the aggregate level disguises the underlying cointegrating relationship that might be otherwise present at a more refined geographic level. The findings suggest that, only at a refined geographic level, the present value relationship seems to be a reasonable constraint to impose on housing prices in the long run. These findings are consistent with various discussions on housing segmentation and show that the long-term cointegrating relationship exists if cross-sectional market heterogeneity has been properly addressed.

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Table 1 Nation-wide Home Purchase-Rental Price Index Pair: 2000q1-2014q4 – Unit Root Test

With an intercept				
	ADF(1)	ADF(2)	ADF(3)	ADF(4)
Log(real purchase price index)	-2.245	-1.428	-1.418	-1.211
ΔLog(real purchase price index)	-4.144***	-3.612***	-3.579***	-4.014***
Log(real rental price index)	-2.572*	-2.403	-1.606	-1.802
ΔLog(real rental price index)	-2.650*	-3.695***	-2.899**	-3.295**
With an intercept and a linear trend				
	ADF(1)	ADF(2)	ADF(3)	ADF(4)
Log(real purchase price index)	-3.615**	-2.502	-2.792	-2.502
ΔLog(real purchase price index)	-4.123***	-3.642**	-3.571**	-4.009**
Log(real rental price index)	-3.104	-2.815	-1.963	-2.320
ΔLog(real rental price index)	-2.590	-3.626**	-2.846	-3.258*

Note: * denotes that the test is significant at the 10% significance level. ** denotes that the test is significant at the 5% significance level. *** denotes that the test is significant at the 1% significance level.

Table 2 Nation-wide Home Purchase-Rental Price Index Pair: 2000q1-2014q4
– Cointegration Test

	ADF(1)	ADF(2)	ADF(3)	ADF(4)
With an intercept	-1.567	-1.318	-0.950	-1.311
With an intercept and a linear trend	-2.207	-1.889	-1.623	-1.842

Note: * denotes that the test is significant at the 10% significance level. ** denotes that the test is significant at the 5% significance level. *** denotes that the test is significant at the 1% significance level.

Table 3 Area-Specific Home Purchase-Rental Price Index Pairs: 2000q1-2014q4 – CD Test

	CD test	p-value	corr	abs(corr)
Log(real purchase price index)	44.47	0.000	0.856	0.856
Log(real rental price index)	46.77	0.000	0.900	0.900

Table 4 Area-Specific Home Purchase-Rental Price Index Pairs: 2000q1-2014q4 – Pesaran's CIPS Panel Unit Root Test

With an intercept				
	CADF(1)	CADF(2)	CADF(3)	CADF(4)
Log(real purchase price index)	-1.709	-1.526	-1.505	-1.069
ΔLog(real purchase price index)	-5.629***	-4.208 ***	-3.656***	-3.185 ***
Log(real rental price index)	-1.729	-1.630	-1.264	-1.107
ΔLog(real rental price index)	-5.983***	-5.454***	-4.364***	-3.943 ***
With an intercept and a linear trend				
	CADF(1)	CADF(2)	CADF(3)	CADF(4)
Log(real purchase price index)	-2.534	-2.497	-2.531	-2.101
ΔLog(real purchase price index)	-5.807***	-4.349***	-3.655***	-3.182**
Log(real rental price index)	-3.096***	-2.499	-1.981	-1.953
ΔLog(real rental price index)	-6.124***	-5.525***	-4.455***	-4.126***

Note: * denotes that the test is significant at the 10% significance level. ** denotes that the test is significant at the 5% significance level. *** denotes that the test is significant at the 1% significance level.

Table 5 Area-Specific Home Purchase-Rental Price Index Pairs: 2000q1-2014q4 – Cointegrating Vector

	MG	CCEMG	CCEP
Log(real rental price index)	0.9908	1.0741	1.000
	(0.125)	(0.123)	(0.112)
Constant	-0.0830	-0.0193	0.0000
	(0.097)	(0.086)	(0.087)
Adjusted R-squared	0.782	0.955	0.926
Average cross-correlation coefficients	0.614	-0.044	-0.061
CD test statistics	31.907	-2.278	-3.145

Note: Standard errors are reported in parentheses.

Table 6 Area-Specific Home Purchase-Rental Price Index Pairs: 2000q1-2014q4 – Cointegration Test

	CADF(1)	CADF(2)	CADF(3)	CADF(4)
With an intercept	-2.172*	-2.084	-2.031	-1.914
With an intercept and a linear trend	-3.019***	-2.843**	-2.988***	-3.022***

Note: * denotes that the test is significant at the 10% significance level. ** denotes that the test is significant at the 5% significance level. *** denotes that the test is significant at the 1% significance level.

Table 7 Area-Specific Home Purchase-Rental Price Index Pairs: 2000q1-2014q4 – Panel Error Correction Estimates

	MG	CCEMG	CCEP
One period lag of Log(real purchase price index) - Log(real rental price index)	-0.1102	-0.1655	-0.1681
	(0.023)	(0.025)	(0.039)
One period lag of ΔLog(real purchase price index)	0.0421	-0.2831	-0.3412
	(0.064)	(0.050)	(0.050)
ΔLog(real rental price index)	0.6308	0.1084	0.0885
	(0.055)	(0.056)	(0.062)
Half life	5.937	3.831	3.766
Adjusted R-squared	0.396	0.715	0.715
Average cross-correlation coefficients	0.315	-0.097	-0.106
CD test statistics	16.071	-4.978	-5.430

Note: Standard errors are reported in parentheses.

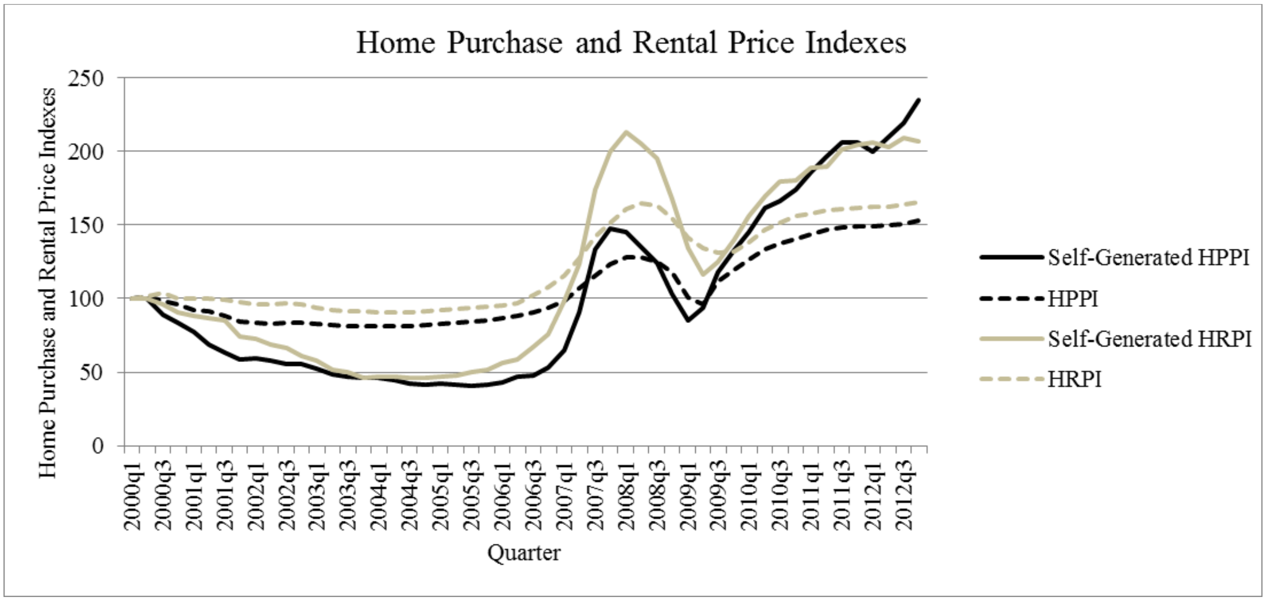

Figure 1: Nation-Wide Home Purchase and Rental Price Indexes

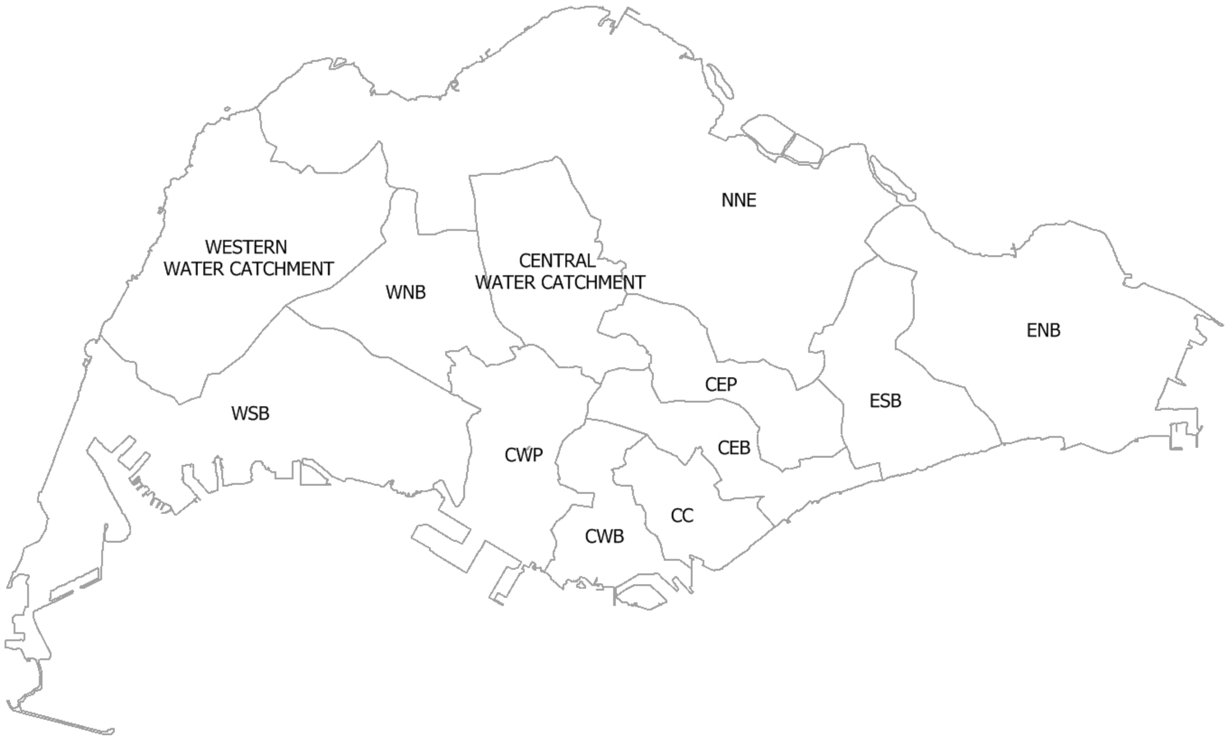

Figure 2: Singapore Urban Planning Area Clusters

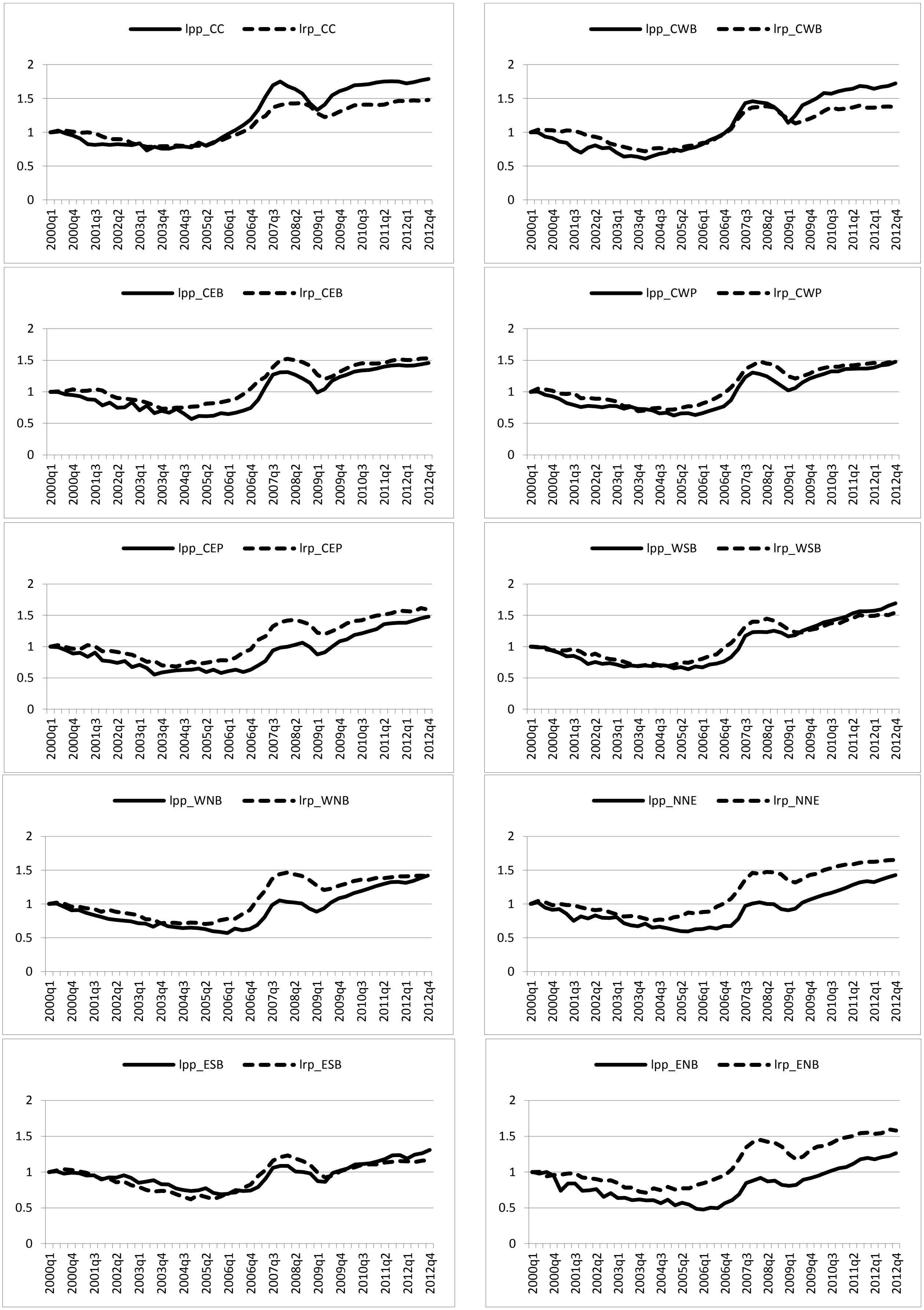

Figure 3: Area-Specific Home Purchase and Rental Price Indexes

Notes: lpp stands for “log purchase price”; lrp stands for “log rental price”; Attached endings in the label correspond to areas as shown in Figure 2.