5 Spatial interrelations of Chinese housing markets: Spatial causality, convergence and diffusion

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Abstract: This paper comprehensively tests the spatial interrelationships of 10 housing markets in the Pan-Pearl River Delta (Pan-PRD) in China, including the properties of spatial causality, convergence and diffusion patterns. The pairwise Toda-Yamamoto Granger causality tests suggest widely existing leading-lag relationships between housing markets; a unidirectional causal flow from the eastern-central area to western China can be tentatively confirmed. However, there is a lack of sufficient evidence supporting pairwise long-run *cointegration* and *convergence*, indicating a diverged interurban housing market in the Pan-PRD. In the short run, the spatial-temporal diffusion model manifests the importance of the spillover effect from neighbouring cities in predicting one city's house price changes. Furthermore, the generalized impulse response functions (GIRFs) clearly depict the transmission pattern of shocks to one chosen city. The diffusion pattern is characterized by the fact that the shocks first spread to nearby cities with cities further away taking a longer time to respond.

Keywords: Spatial causality, long-run convergence, ripple effect, diffusion, house prices, China

JEL: C31, C33, R21, R31

§ 5.1 Introduction

After the financial crisis in 2008, many governments have been attempting to stimulate depressed housing markets through policy interventions. However, whether the interventions can work as expected relies heavily on our understanding of housing markets. To provide deeper insights into house price behaviour, many scholars advocate an investigation into a series of interrelated regional markets rather than a single national market (Meen 1996; Yunus and Swanson 2013). Indeed, the structure of regional housing markets is likely to vary significantly across space given the huge differences in local amenities, economic conditions, and regulation constraints, among other considerations. Simply aggregating a bundle of local housing markets into a national unit could lead to severe misunderstanding, particularly in a country in which the regional housing markets are highly differentiated.

Regional housing markets are neither identical nor independent. A large volume of literature has provided evidence supporting the interrelations of regional housing markets (e.g.,Giussani and Hadjimatheou 1991; Pollakowski and Ray 1997; Holly et al. 2011). Specifically, researchers find that house price changes in an area depend significantly on what occurred in other areas' housing markets. Among the various interrelations of local housing markets, long-run *integration*, which describes a situation in which local house prices maintain an equilibrium relation in the long-run, has long been a concern because of its policy implications. If local housing markets are highly integrated, a unified nation-wide housing policy will be sufficient; otherwise a basket of diversified, locally-oriented policies are necessary. Another parallel research agenda has concentrated on the so-called ripple effect whereby house price shocks to an area will gradually diffuse to other areas, with areas further away being slower to respond to the shocks. Statistical evidence for long-run integration and a ripple effect of regional house prices has been found, for example, in the UK markets by Alexander and Barrow (1994), Meen (1996), Cook (2003) and Holmes and Grimes (2008), although certain studies cast doubt on it (Drake 1995; Ashworth and Park 1997; Abbott and Vita 2013) [1].

While a large amount of empirical evidence for long-run and short-run patterns of house prices is already available, the underlying behavioural mechanisms are not yet clear. Meen (1999) proposed five possible explanations for the patterns in the UK market: migration, equity transfer, spatial patterns in the determinants of house prices, spatial arbitrage, and coefficient heterogeneity. Although the transitional economy of China makes its housing market significantly different from the UK market, we have observed the presence of such factors which can cause a certain pattern of house prices. For example, the loosening of *Hukou* restrictions has largely accelerated labour mobility between areas and consequently induced the equity transfer among regions [2]. The information transmission pattern, namely, that housing market information usually flows from "superstar" cities to "normal" cities (Wu and Deng 2015), raises the chance of spatial arbitrage. Hanink et al.(2012) showed significant

1 It should be noted that the mixture of the evidence is partly due to the different understanding of the term '*integration*' ('*convergence*') and '*ripple effect*'. We will discuss this in the literature review.

2 The "*Hukou*" (household registration) system in China was initially designed as a mechanism of monitoring population movements in early 1950s. Afterwards, it became a strong tool to restrain the rural-urban migration and the labour mobility between cities. Since 1980s, the power of "*Hukou*" system has been weakened through a series of reforms, but it remains in place to this day.

coefficient heterogeneity among Chinese county-level housing markets. However, whether these factors can result in long-run integration of regional housing markets remains unclear. From the national perspective, the current migration pattern, flowing from less developed Western China to developed Eastern China, is likely to induce the divergence of housing markets between the East and West rather than market integration. However, local housing markets within the East or the West have a larger chance to be integrated. Spatial patterns of house price determinants also provide us with a confusing hint regarding the long-run integration of local house prices. Province-level real GDP per capita, used as a proxy for income, is found to be convergent in Eastern and Western China, but not in Central and North-eastern China (Su and Chang 2013).

Given such arguments, the spatial interrelations of Chinese local housing markets appear to be an interesting question to answer. Indeed, much effort has been dedicated to this issue in recent years. For example, Wang et al. (2008) examined the long-run and short-run properties of house prices based on cities within 5 sub-national areas during the period 1997Q4 – 2007Q1. Huang et al. (2010b) conducted research on nine major Chinese cities during a similar time span (1999Q1 – 2008Q3), and Li and Li (2011) on nine cities in Pearl River Delta for the period 2001Q1 – 2010Q4. In general, these studies confirmed the spatial interrelations of housing markets among different cities and they found long-run equilibrium relationships between these markets.

Using a new data set of house price indexes for 10 cities within the Pan-Pearl River Delta (Pan-PRD) spanning from June 2005 to May 2015, this paper comprehensively investigates the spatial-temporal interrelations between city-level housing markets. Specifically, we are particularly interested in the following three questions. First, is there any 'spatial causality' in the interurban housing markets so that the historical house price information in one market can be used to predict the current house prices in other markets? Second, are the house price indexes of ten cities converged (integrated) or segmented in the long-run? Third, is there a distinct house price diffusion pattern so that shocks to one particular market can propagate to other markets gradually?

To our knowledge, this is the first paper that focuses on the spatial interrelations of housing markets in the Pan-PRD area in South China. This area is of great interest given its economic importance and policy implications. Since the reform and opening-up started at 1978, the cities of Pearl River Delta (PRD) in Guangdong province, such as Shenzhen and Guangzhou, have been rapidly developing due to their advantageous location and access to Hong Kong and Macao. Meanwhile, most Central and Western provinces, which provide a large amount of cheap labour for Guangdong and thus can be seen as the hinterland, still struggled with low economic growth. To narrow the gap of development between these areas, "Pan-Pearl River Delta Regional

Co-operation Framework Agreement" was signed by 11 relevant governments in 2004. This initiative aims to remove the trade barriers between cites, promote the economic linkages and interaction between eastern, central and western China, and finally achieve the economic integration of this area. The results of this paper shed light on the extent to which the cities in this cooperation framework are linked with each other and the degree to which their markets have been integrated. Thus, this paper might have great implications for policy makers.

Our results suggest widely existing pairwise leading-lag relations among the housing markets under investigation. That is, a city's housing market is generally interrelated with the markets of other cities. However, in contrast to most of the previous studies that support the long-run integration of interurban housing markets, we find rare evidence for pairwise cointegration relationships between cities in the Pan-PRD, and even less evidence for convergence. This discrepancy is probably due to the fact that we focus on a large and heterogeneous area, while previous studies are confined to a relatively small and homogeneous area or to the Chinese cities that have similar socio-economic conditions. Furthermore, a distinct house price diffusion pattern is confirmed; the generalized impulse response function (GIRF) shows that shocks to a city first spread to the nearby cities and then gradually to the distant cities.

The remainder of this paper is organised as follows. Section 2 briefly reviews the related literature, followed by the data description in section 3. The empirical examination of the leading-lag relationships, long-run integration and house price dynamic pattern are shown in section 4, 5 and 6, respectively. Finally, section 7 concludes the findings and derives certain implications.

§ 5.2 Previous literature

The focus on regional housing markets interaction dates to the observation of UK housing markets: house price disparities between South and North tended to increase in the 1980s, but tended to narrow again in the 1990s (Giussani and Hadjimatheou 1991). This behaviour inspires the discussion on regional market integration and the 'ripple effect' hypothesis.

The long-run properties of regional house prices are usually examined under the cointegration framework. MacDonald and Taylor (1993) and Alexander and Barrow (1994) found general evidence for cointegration relationships between regions within either the South or the North of the UK, although the South/North segmentation still appears to exist. In the U.S. housing markets, Yunus and Swanson (2013) documented systematic cointegration among 9 census regions, the degree of which has further intensified after the subprime crisis.

Certain researchers take the idea of cointegration one step further and are interested in the long-run *convergence*, which describes a situation in which local house prices converge towards a constant equilibrium relationship in the long-run (Meen 1996) [3]; that is a more stringent concept than *cointegration*. Cointegration is necessary but not sufficient for long-run convergence of regional markets. House price convergence necessitates that two house price series are cointegrated with a cointegrating vector following the form (1,-1), as well as that they are co-trending, which means no deterministic trend in the cointegrating vector (Holly et al. 2011; Abbott and Vita 2013). In accordance with this tradition, Meen (1996) tentatively suggested three groups (namely the South, the North and the Midlands) in the UK within which house prices may be converged. However, a later study by Abbott and Vita (2013), using the pairwise approach, offered no evidence in support of overall convergence or 'club convergence' [4]. Controversially, Holmes et al. (2011) applied the pairwise approach to the US market and found significant supportive evidence of long-run convergence between state house prices, as well as between MSA house prices.

Since Meen (1999) noted that long-run convergence is equivalent to the long-run stationarity of deviations of regional house prices from the national average, another strand of studies uses the unit root test of the ratio of regional/national house prices to investigate long-run convergence properties. Although Meen (1999) failed to prove the stationarity of region/national ratios using an augmented Dickey-Fuller (ADF) test, Cook (2003) successfully reversed the negative findings by using threshold autoregressive (TAR) and momentum threshold autoregressive (MTAR) tests, which can allow for asymmetric adjustments. The researcher contended that the failure of previous studies is due to the neglect of significant asymmetry in the convergence process. More recently, the stationarity of regional/ national house price ratios in the UK market was confirmed by Holmes and Grimes (2008) who conducted the unit root test on the first principal component (FPC). By applying non-linear unit root tests and linear unit root tests with structural breaks, Canarella et al. (2012) documented conflicting evidence in favour of the stationarity of U.S. metropolitan house price indices to a national house price index [5].

3 The convergence here is commonly referred to as stochastic convergence from the time-series point of view. It does not imply that all the local house prices are equalized across regions. However, another notion of convergence that house prices will ultimately converge to the common level in the long run is also investigated by, for example, Kim and Rous (2012) and Blanco et al. (2016).

4 The pairwise approach is essentially similar with Engle-Granger two-step cointegration procedure (Engle and Granger 1987), but pre-specifies the cointegrating coefficients to be (1,-1) in the first step. In this case, the normal unit root statistic can be used to test the unit root of cointegrating residuals in the second step. If the null hypothesis of unit root and linear trend are rejected, the house prices are considered to be converged.

5 In addition to the methods we noted, the battery of approaches that have been dedicated to the long-run convergence of house prices also includes the Kalman filter/time varying parameter (TVP) estimation (Drake 1995) and the so-called synchronicity approach (Miles 2015).

In the UK market, house price changes are usually first observed in London, and then spread to other regions, with the distant regions echoed last. This behaviour is often referred to as a 'ripple effect'. This phenomenon is proved by Meen (1996) who revealed that the speed of adjusting to an equilibrium relation with the South East for each region clearly falls as moving to the North. However, Ashworth and Park (1997) suspected this assumption because they found that the other regions have the common timing of responses to shocks from the South East. Generally, a ripple effect could mean the propagation of shocks emanating from a 'dominant' market such as London to the remaining markets, ignoring the response time of each market. Thus, MacDonald and Taylor (1993) intuitively exhibited a ripple effect from Great London to other regions by using impulse response functions. Alexander and Barrow (1994), using the Granger causality test, detected a causal flow from south to the north with the South East being the most likely exogenous region.

Compared with the term 'ripple effect', U.S. researchers appear to prefer the term 'spatial diffusion', which emphasises the influence on one specific housing market originating from neighbouring markets, not solely from a certain 'dominant' market. An example is Pollakowski and Ray (1997) who revealed the importance of an area's historic price change information in predicting other areas' price change at both the primary metropolitan statistical area (PMSA) level and the subnational census division level.

A recent development in modelling house price diffusion is to incorporate the spatial correlations of housing markets into the conventional time-series models. For example, Holly et al. (2011) proposed a spatial diffusion model in which the house price changes of a region are affected by the short-term and the long-run house price changes both in London and in neighbouring regions. Additionally, the spillover effect of neighbouring regions is evident in the estimation results. Similarly, when modelling U.S. state house prices, Brady (2014) adopted a so-called single-equation spatial autoregressive panel model, which incorporates a "spatial regressor" that is common to spatial autoregressive models, into the panel model. The spatial impulse response functions (SIRFs) support the existence of spatial diffusion.

The relevant studies on long-run and short-run properties of local housing markets in the literature are not exhausted. Certain other examples include Stevenson (2004) on the market of the Republic of Ireland, Berg (2002) on the Swedish second-hand market, Balcilar et al. (2013) on the 5 major metropolitan area markets of South Africa, and Luo et al. (2007) on state capital cities in Australia. More recently, a few studies focusing on Asian housing markets have emerged. For example, Lean and Smyth (2013) documented a ripple effect from the most developed states to the less developed states of Malaysia. A ripple effect from the central city to the suburbs is also demonstrated in Singapore (Liao et al. 2015). In Taiwan, both Lee and Chien (2011) and Chen et al. (2011) offered partial evidence for long-run stable relationships across an inter-city housing market; however, Taipei, the economic centre and capital city, appears not to be the Granger causality of regions in Southern Taiwan.

For historical reasons, urban private housing markets in China were not established until the late 1990s. The lack of continuous house price records makes it difficult to investigate the interaction of local housing markets; however, few attempts have been made by scholars in mainland China. Wang et al. (2008) examined spatial interrelations of house prices among the cities within 5 sub-national areas during the period 1997Q4 – 2007Q1 [6]. Within each of the five sub-national markets, Johansen cointegration test suggests the existence of at least two cointegration relationships; moreover, they found heterogeneous diffusion patterns. Meanwhile, Huang et al. (2010b), focusing on the pair-wise relationships among nine major Chinese cities during a similar time span (1999Q1 – 2008Q3), also presented evidence for generally existing long-run equilibrium relationships. Li and Li (2011) found cointegration relationships for the 9 cities in Pearl River Delta for the period 2001Q1 – 2010Q4; furthermore three submarkets are identified based on Granger Causality test. In addition, Huang et al. (2010a) used a two-stage procedure to test the ripple effect hypothesis in 19 cities based on a period from January 2008 to April 2010. The ripple effect is supported by the evidence that popular and vibrant cities that have greater price fluctuations, such as Guangzhou and Shenzhen, are likely to achieve a turning point earlier than other less active cities.

This paper attempts to offer a comprehensive investigation into the interrelations of Chinese housing markets in an economic co-operation framework – the Pan-PRD, which has not yet been considered by previous studies. First, we examine whether housing markets depend on each other through a Granger causality test. Second, the pairwise long-run *cointegration* and *convergence* properties are tested, respectively. Finally, when modelling the house price diffusion pattern, we consider spatial dependence in accordance with the treatment by Holly et al. (2011).

§ 5.3 Data

§ 5.3.1 The “Pan-Pearl River Delta”

This paper focuses on the housing market interrelations among prefecture-cities [7]. The

6 The five sub-national areas are Northern Coast Area, Central Coast Area, Southern Coast Area, Central Area and Western Area.

7 A prefecture city is an administrative division of China, ranking in the second level of administrative structure. Under its administration are the counties (county-level cities) and districts, of which the districts constitute the city proper. The housing market we noted in this paper pertains to the *city proper*.

whole area under investigation is the so-called "Pan-Pearl River Delta" (Pan-PRD) in South China. The Pan-PRD is a regional co-operation framework launched in June 2004, which has the objective to remove the barriers to the flow of production factors and finally establish a common market. This framework is composed of 11 geographically contiguous spatial units, including 9 provinces: Guangdong, Fujian, Jiangxi, Hunan, Guangxi, Guizhou, Yunnan, Sichuan and Hainan; plus 2 special administrative regions: Hong Kong and Macao (known as "9+2") (refer to Figure 5.1).

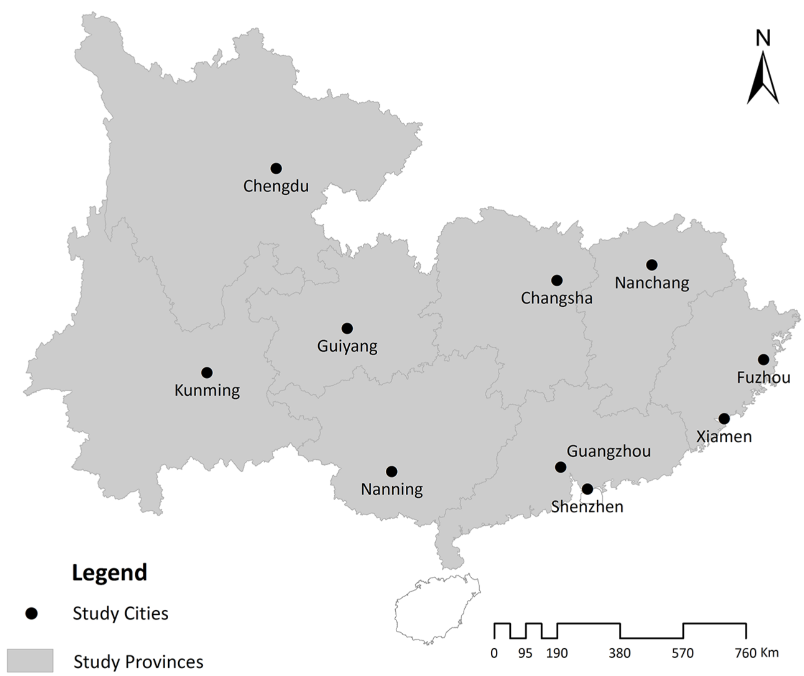


FIGURE 5.1 "Pan-Pearl River Delta" and study cities

The Pan-PRD is already the largest economic bloc in China, representing 20% of China's total land area, 36% of its population and 40% of its GDP (2004 figure). The Pan-PRD spans across several geographic and economic zones that are formulated by the central government. Guangdong, Fujian, Hong Kong and Macao, bordering the South China Sea, are categorised as part of Eastern China; Hunan and Jiangxi, connected with the Yangtze River, belong to Central China, and the remaining four inland provinces are divided into Western China. Obviously, economic development in the Pan-PRD area is far from integration, and regional disparities remain notable. In general, aside from Hong Kong and Macao, the eastern provinces, Guangdong and Fujian, are much more developed than the remaining inland provinces. In particular, Guangdong is undoubtedly the leading province due to its production capabilities and its economic integration with Hong Kong.

The 10 cities of interest in the Pan-PRD are the capitals of 8 provinces, Guangzhou (Guangdong), Fuzhou (Fujian), Nanchang (Jiangxi), Changsha (Hunan), Nanning (Guangxi), Guiyang (Guizhou), Kunming (Yunnan) and Chengdu (Sichuan), and 2 special economic zones (SEZs), Shenzhen (Guangdong) and Xiamen (Fujian). Haikou, the capital of Hainan province, is excluded because its house price development path is clearly different from the others [8]. With increasing economic cooperation in the Pan-PRD, we believe that the dependence between the 10 housing markets is also strengthened.

§ 5.3.2 House price index

The availability of the house price data has been the largest obstacle to examining house price behaviour on temporal and spatial dimensions. Because a truly private housing market was not developed in most Chinese cities until 1998, house price indices used for measuring the movement of house prices are rare. "Price Indices for Real Estate in 35/70 Large- and Medium-sized Cities" (70 Cities Index), compiled and published by the National Bureau of Statistics of China (NBSC), is the only public accessible index system that can provide consistent information on house prices over a long period.

The 70 Cities Index system was first established in 1997 and originally covered 35 major cities. The system published year-over-year quarterly price changes for land transactions, housing sales and housing rentals. However, with the rapid growth of housing transactions and house prices after 2000, the index system could no longer accurately reflect house price movement and was widely criticised by the public. Therefore, NBSC updated the survey and calculation method, which can be called the "matching approach" (Wu et al. 2014), to obtain the quality-adjusted price index, and the scope was expanded to 70 large- and medium-sized cities. Since July 2005, the 70 Cities Index system has been formally reported on a monthly basis, including year-over-year and month-over-month Laspayres indices reported monthly. In January 2011, the survey and calculation strategy for the 70 Cities Index was refined again. Since then, the chained Laspayres index was also made available (base year of 2010).

We apply the "Price Indices of Newly Constructed Residential Buildings" (NCRB Index), drawn from the 70 Cities Index system, for the 10 cities in our empirical analysis,

8 As the capital of Hainan, the sole tropical island in China, Haikou is a famous tourist city. Home buyers from outside constitute a very large share of housing demand so that is no surprise that the housing market of Haikou has certain distinctive characteristics and differs from others.

covering the period from June 2005 to May 2015, for a total of 120 observations. Quarterly year-over-year indexes prior to 2005 are discarded from the analysis. First, the index in this period was likely unreliable due to the rough survey and the calculation methodology. Second, during the period after 2005, house prices are very volatile due to active housing market transactions. Thus, monthly data are more appropriate to model this volatility in house prices.

For our analysis, we joined together the June 2005 – December 2010 index and the January 2011 – May 2015 index. When compiling the house price index by chaining the month-over-month house price changes, at first, January 2011 was set as a reference base because the NCRB Index did not report consistently before and after 2011. Since the month-over-month index before 2010 is not transitive, there is a drift in the early years of house price index [9]. Thereafter, the index series was re-based to June 2005.

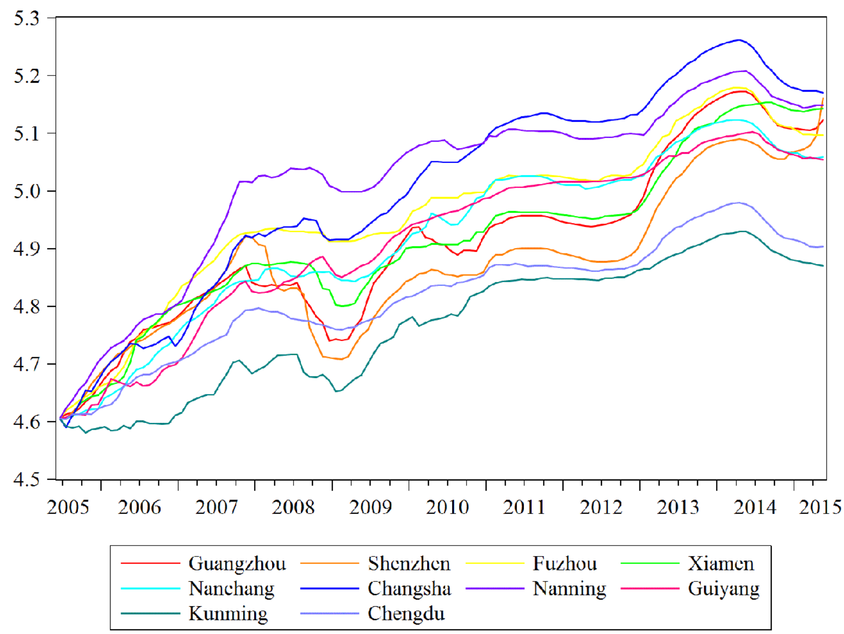


FIGURE 5.2 "Pan-Pearl River Delta" and study cities

Figure 5.2 shows the natural logarithm of house price indices for 10 cities. A casual examination of the figure suggests that house price developments of the 10 cities

9 We note that the chained month-over-month index is not identical to the chained year-over-year index. However, the house price series obtained from chaining these two indexes are highly correlated (the correlation coefficient is greater than 0.97). There are only small differences for the period before January 2007.

follow a similar upward trend throughout the study period. The house prices of Guangzhou and Shenzhen may be much more volatile than those of other cities, particularly during the period after the 2008 global financial crisis, which experienced a sharp decrease of house prices.

§ 5.4 Spatial leading-lag relationships

§ 5.4.1 Toda-Yamamoto Granger causality test

The first question of whether there is a spatial leading-lag relationship or a causal flow between intercity housing markets can be examined using Granger causality tests. Suppose two house price series, p_{1t} and p_{2t}, take the form of VAR(k)

$$\begin{aligned} p_{1t} = c_{10} + & a_{11}^{1} p_{1t-1} + a_{11}^{2} p_{1t-2} + \cdots + a_{11}^{k} p_{1t-k} \\ & + a_{12}^{1} p_{2t-1} + a_{12}^{2} p_{2t-2} + \cdots + a_{12}^{k} p_{1t-k} + \epsilon_{1t} \end{aligned} \quad (1)$$

$$\begin{aligned} p_{2t} = c_{20} + & a_{21}^{1} p_{1t-1} + a_{21}^{2} p_{1t-2} + \cdots + a_{21}^{k} p_{1t-k} \\ & + a_{22}^{1} p_{2t-1} + a_{22}^{2} p_{2t-2} + \cdots + a_{22}^{k} p_{2t-k} + \epsilon_{2t} \end{aligned} \quad (2)$$

The series p_{2t} is said to Granger cause p_{1t} if the historical values of p_{2t} can contribute to predicting p_{1t} in equation (1). This is equivalent to a test of the null hypothesis $H_0 : a_{12}^{1} = a_{12}^{2} = \cdots = a_{12}^{k} = 0$. The standard causality test procedure requires that the series p_{1t} and p_{2t} are stationary. If they are both $I(1)$, a pre-test of cointegration is needed. In this paper, we act in accordance with the Toda and Yamamoto (1995) (TY) procedure, which overcomes the limitations of standard methodology in a manner that can allow the series to be integrated or cointegrated of an arbitrary order. In other words, the TY procedure estimates an augmented VAR(k+d) system in which d is the maximum order of integration of the time series in the system. The Granger causality tests are then performed on the first k coefficient matrices by using a standard Wald test, ignoring the coefficients matrices of the last d lagged vectors.

§ 5.4.2 Results

We begin by determining the integration orders of 10 (log) house price series. The commonly used Augmented Dickey-Fuller test (Dickey and Fuller 1979) is performed, and the results are reported in the first and fourth columns of Table 5.1. The null hypothesis of unit root for *level* variables of Guangzhou and Xiamen is rejected at the approximately 5% significance level (the p-value of Xiamen is 5.76%), indicating a trend stationary process for these two cities. Conversely, all the *first difference* series are stationary. That is, house price indexes of 8 of 10 cities are unit root processes and are integrated of order one, i.e., $I(1)$. The ADF tests provide preliminary evidence that house prices of different cities are integrated of different orders.

The ADF test has very low power in distinguishing highly persistent stationary processes from non-stationary processes; the power is lower when a deterministic trend is included in the test. Thus, an efficient unit root test, the Dickey-Fuller generalized least square (DF-GLS) test proposed by Elliott et al. (1996), is also conducted. The results in the second and fifth column of Table 5.1 are largely consistent with the ADF test, except that the level house price determination of Xiamen is a unit root process.

TABLE 5.1 Unit root test

	level			*1st difference*		
	ADF	DF-GLS	KPSS	ADF	DF-GLS	KPSS
Guangzhou	-3.886(2)**	-3.361(2)**	0.080(9)	-3.467(1)**	-3.494(1)**	0.056(8)
Shenzhen	-2.179(1)	-2.209(1)	0.121(9)*	-3.477(0)**	-3.510(0)**	0.078(8)
Fuzhou	-2.538(2)	-1.399(2)	0.157(9)**	-3.642(1)**	-3.574(1)***	0.085(8)
Xiamen	-3.391(2)*	-2.404(2)	0.103(9)	-3.784(1)**	-3.642(1)***	0.086(8)
Nanchang	-1.411(1)	-0.957(1)	0.199(9)**	-6.037(0)***	-5.233(0)***	0.051(8)
Changsha	-1.852(1)	-0.962(1)	0.226(9)***	-6.596(0)***	-3.899(0)***	0.046(8)
Nanning	-2.188(1)	-0.862(1)	0.248(9)***	-4.579(0)***	-4.382(0)***	0.083(8)
Guiyang	-1.044(1)	-0.745(1)	0.286(9)***	-6.289(0)***	-5.908(0)***	0.042(8)
Kunming	-1.365(1)	-1.764(1)	0.148(8)**	-7.070(0)***	-5.161(0)***	0.092(7)
Chengdu	-2.623(2)	-2.144(2)	0.127(9)*	-3.628(1)**	-3.347(1)**	0.055(8)

Notes: All the models include an intercept and a deterministic trend. Numbers shown in parentheses are the lag length or bandwidth. For the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979), the lag length for *level* variables is selected using the Bayesian Information Criterion (BIC) with the maximum lag length being set to 12. DF-GLS unit root tests (Elliott et al. 1996) use the same lag length chosen for ADF test. For KPSS test (Kwiatkowski et al. 1992), the bandwidth is selected by Newey-West method. In all the cases, the lag length for *first difference* variables equals to the lag length for *level* variables minus one. The null hypothesis for ADF and DF-GLS tests is having a unit root, but stationary for the KPSS test. *, ** and *** indicate significance at the 10%, 5% and 1% level, respectively.

Table 5.1(column 3 and 6) also presents the results of the KPSS test (Kwiatkowski et al. 1992), which has the null hypothesis of stationary. The KPSS results support that Xiamen's house price series is $I(0)$, in accordance with the ADF test but in contradiction to the DF-GLS test. The evidence for the integration order of the Xiamen price series appears to be mildly confusing. Given that the KPSS statistic is very close to the 10% critical value (0.119), it is reasonable to assume Xiamen's price process to be I(1) for the following analysis [10].

10 One reviewer pointed out that house prices are very likely to display an asymmetric adjustment, reducing the

Because not all the house price series are integrated of the same order, the pairwise Toda-Yamamoto Granger causality procedure is preferable in this analysis. According to information criteria, such as AIC and BIC, the optimal lag lengths for most of the city-pair VARs are 2 and 3. To largely ensure that the residuals are close to white noise, we proceed using the VAR($3 + d$) system in which d is the maximum integration order of the city pair.

The results of Toda-Yamamoto Granger causality tests are reported in Table 5.2, where the null hypothesis is that column cities do not Granger cause row cities. Overall, in most of the city pairs, we find significant evidence of bilateral or unidirectional leading-lag relations; however, the causality pattern is very complicated [11]. Certain literature focusing on housing market interaction, such as Clapp et al. (1995) and Chen et al. (2011), has claimed that the house price interrelation (or causality) only occurs between neighbouring markets. Our results cast doubt on this conclusion because leading-lag relationships are found in many city pairs where the two cities are separated by very long distances. Conversely, there is no spatial causality in certain short-distance city pairs, such as Xiamen and Fuzhou. Similarly, Pollakowski and Ray (1997) and Luo et al. (2007) also found significant causality between non-contiguous regions. Such complicated causality patterns may be largely due to economic relations rather than behavioural reasons (Pollakowski and Ray 1997).

To further examine the results, the cities in the system are divided into two groups: the eastern-central group including the first six cities in Table 5.2 and the western group containing the remaining cities. We tentatively find a general unidirectional causal flow from the eastern-central area to western China. The historical house price information of all eastern-central cities, except Xiamen, can significantly contribute to predicting the house prices of western cities. The opposite, conversely, can hardly be true given the largely insignificant Granger causality test results in the lower-left panel. However, among the four western cities, the housing markets of Kunming and Chengdu appear to play a role in predicting house price behaviours in eastern-central cities. The above findings are closely related to the socio-economic disparities between eastern, central and western China. Considering that eastern-central cities are generally more developed than western ones, it can be expected that their market dynamics can lead housing market behaviours in the remaining cities. Chengdu is an exception in western cities, given its status as the financial and economic centre of western China. Consequently, the mutual leading-lag relationships between Chengdu and most

power of traditional unit root tests. Thus, we employed the momentum threshold autoregressive (MTAR) asymmetric unit root test proposed by Enders and Granger (1998) to test the integration order of 10 cities' house price indexes. However, the results, which are available upon request, do not turn over the finding that all the cities are $I(1)$ process except for Guangzhou.

11 It should be noted that, throughout the paper, when we say "a market leads or causes another market in the Granger sense", we cannot exclude the possibility that such correlation is caused by common shocks.

TABLE 5.2 Toda-Yamamoto Granger causality test

	Guangzhou	Shenzhen	Fuzhou	Xiamen	Nanchang	Changsha	Nanning	Guiyang	Kunming	Chengdu
Guangzhou		25.63***	6.47*	50.54***	11.13***	14.10***	11.25***	9.47**	4.41	9.16**
		(0.000)	(0.091)	(0.000)	(0.011)	(0.003)	(0.010)	(0.024)	(0.220)	(0.027)
Shenzhen	13.82		9.14**	18.93***	6.05*	18.76***	8.25**	14.14***	6.22*	8.23**
	(0.003)		(0.027)	(0.000)	(0.109)	(0.000)	(0.041)	(0.003)	(0.101)	(0.041)
Fuzhou	2.00	1.52		7.52*	5.04	7.97**	8.52**	6.49*	8.26**	6.78*
	(0.572)	(0.678)		(0.057)	(0.169)	(0.047)	(0.036)	(0.090)	(0.041)	(0.079)
Xiamen	18.73***	2.32	3.38		4.10	2.82	5.28	2.01	0.28	0.94
	(0.000)	(0.510)	(0.336)		(0.251)	(0.420)	(0.153)	(0.570)	(0.963)	(0.815)
Nanchang	4.47	7.90**	8.25**	6.42*		14.61***	12.22***	17.52***	11.44***	11.55***
	(0.215)	(0.048)	(0.041)	(0.093)		(0.002)	(0.007)	(0.001)	(0.010)	(0.009)
Changsha	1.03	5.31	7.48*	6.11*	6.74*		13.36***	34.22***	13.29***	7.99**
	(0.794)	(0.150)	(0.058)	(0.106)	(0.081)		(0.004)	(0.000)	(0.004)	(0.046)
Nanning	8.91**	14.22***	4.92	3.52	5.08	9.51**		19.42***	8.64**	8.70**
	(0.031)	(0.003)	(0.178)	(0.318)	(0.166)	(0.023)		(0.000)	(0.035)	(0.034)
Guiyang	0.41	2.37	2.32	0.67	5.31	5.07	5.29		1.38	7.46*
	(0.938)	(0.500)	(0.509)	(0.880)	(0.150)	(0.167)	(0.152)		(0.711)	(0.058)
Kunming	7.27*	11.59***	1.79	5.96	7.30*	27.03***	8.31**	11.68***		10.32**
	(0.064)	(0.009)	(0.616)	(0.113)	(0.063)	(0.000)	(0.040)	(0.009)		(0.016)
Chengdu	5.57	5.24	16.81***	17.69***	15.25***	8.21**	6.20*	7.24*	7.19*	
	(0.135)	(0.155)	(0.001)	(0.001)	(0.002)	(0.042)	(0.102)	(0.064)	(0.066)	

Notes: Toda-Yamamoto Granger causality test is performed under VAR(3+d) system where d is the maximum integration order of house price processes in the city pair. Both intercept and deterministic time trend are included in the VAR system. The null hypotheses are that house prices in column cities do not Granger cause house prices in row cities. The asymptotic Wald statistics are reported, and the corresponding p-values are shown in parentheses. *, ** and *** indicate significance at the 10%, 5% and 1% level, respectively.

eastern-central cities are no surprise. Furthermore, within either the eastern-central or western group, the significant leading-lag relationship (at least for one direction) can be found in every city pair.

Given the leading position of Guangdong in the Pan-PRD, we expect that the housing markets of the two cities under its territory, namely Guangzhou and Shenzhen, will lead the markets in other cities. Indeed, Guangzhou and Shenzhen impose a significant leading influence on nearly all the other housing markets, and they are less predictable on the basis of previous information from other markets. However, these cities' dominant role is not unique. Certain other cities appear to also have similar 'exogenous' properties, such as Nanchang.

§ 5.5 Long-run properties

§ 5.5.1 Cointegration and convergence test

The previous section reveals that the 10 cities' housing markets in the Pan-PRD are interrelated with each other. In this section, we go further to ask the question of whether these markets are tied together in the long-run, i.e., if they hold a long-run equilibrium relationship. To answer this question, the long-run cointegration and convergence properties of house prices are investigated.Two $I(1)$ house price series p_{1t} and p_{2t} are said to be cointegrated if a linear combination of p_{1t} and p_{2t} is stationary. Since we are interested in the pairwise cointegration of house price indexes of two cities, the Engle-Granger (EG) two-step procedure (Engle and Granger 1987) is employed in this paper, which has been applied by, for example, MacDonald and Taylor (1993) to a similar question.

The first step is to estimate the long-run equilibrium relationship by the following equation

$$p_{it} = D + \beta p_{2t} + u_t \tag{3}$$

where D is deterministic terms that may contain a constant, a deterministic trend or both. The cointegration test involved in the second step is then based on testing the unit root of residual series u_t. If u_t is stationary, we say that p_{1t} and p_{2t} are cointegrated with $(1, -\beta)$. Because of the spurious regression under the null hypothesis of non-cointegration in the first stage, the residual-based ADF test in the second stage does not have the standard Dickey-Fuller distribution. Therefore, critical values for cointegration test simulated by MacKinnon (1996) are used in this paper.

As noted by Holly et al.(2011) and Abbott and Vita (2013), conditional on cointegration, the long-run convergence of house prices necessitates two additional conditions: (1) the cointegrating vector corresponding to house price series being (1,-1), and (2) the lack of a deterministic trend being presented in cointegration space. The long-run convergence property is tested by the so-called pair-wise approach, which has been used by Holmes et al. (2011) and Abbott and Vita (2013). Specifically, the cointegrating vector is pre-specified with form (1,-1), and then any standard unit root test can be directly used to test the stationarity of cointegrating residuals.

TABLE 5.3 Engle-Granger pairwise cointegration results

	Shenzhen	Fuzhou	Xiamen	Nanchang	Changsha	Nanning	Guiyang	Kunming	Chengdu
Shenzhen		-0.633	-2.053	-0.837	-0.532	-0.520	-0.919	-0.723	-0.067
		-3.094	-13.706	-4.413	-2.092	-1.557	-3.967	-2.718	-0.280
Fuzhou	-0.221		-0.855	-2.094	-1.716	-1.981	-2.204	-2.239	-2.711
	-1.289		-3.683	-5.553	-5.963	-7.077	-9.371	-7.201	-10.044
Xiamen	-1.875	-1.129		-0.791	-1.362	-1.248	-1.482	-1.339	-0.831
	-13.728	-4.692		-2.390	-5.642	-3.800	-5.885	-4.200	-3.591
Nanchang	-0.501	-2.423	-0.604		-3.792*	-2.136	-3.086*	-3.172	-2.615
	-3.266	-6.421	-1.918		-27.395**	-6.507	-19.217*	-14.187	-14.058
Changsha	-0.025	-1.774	-1.279	-3.722*		-1.657	-3.331*	-4.069**	-2.381
	-0.129	-6.234	-5.748	-26.892**		-3.287	-20.505*	-20.546*	-11.213
Nanning	0.309	-1.762	-0.693	-1.693	-1.162		-1.988	-1.938	-1.593
	1.225	-6.234	-2.155	-5.007	-2.189		-6.342	-5.058	-4.286
Guiyang	-0.580	-2.387	-1.498	-3.086*	-3.435*	-2.453		-2.907	-2.680
	-3.231	-10.249	-6.588	-19.214*	-21.148*	-8.041		-14.028	-15.741
Kunming	-0.511	-2.793	-1.438	-3.470*	-4.479**	-2.856	-3.118		-3.169
	-2.369	-8.815	-4.608	-15.207	-22.088*	-7.390	-14.779		-13.768
Chengdu	0.330	-2.919	-0.661	-2.569	-2.420	-2.017	-2.601	-2.843	
	1.649	-10.753	-2.936	-13.714	-11.287	-5.444	-14.893	-12.465	

Notes: The two statistics in each cell are, respectively the τ statistic and z statistic of ADF test with the null hypothesis of no cointegration. The MacKinnon (1996) critical values are used. In each cointegration equation, the row cities are defined as dependent variables. The constant is included in the cointegration space in all city pairs except the Nanchang-Guiyang pair, in which the constant is not significant. * and ** indicate the 5% and 1% significance level, respectively.

§ 5.5.2 Empirical results

Table 5.3 demonstrates the results of the pairwise Engle-Grange cointegration test. Guangzhou is excluded from cointegration analysis because it is $I(0)$ according to our unit root test results. A brief view of the results indicates that cointegration relationships rarely exist between cities, in contrast to the widely existing leading-lag relationships. Among the city pairs that are tied together in the long-run, three cities, Nanchang, Changsha and Guiyang, form a 'cointegration club' within which every city cointegrates with each other. In addition, the significant long-run equilibrium relationship can also be observed in the city pair of Changsha-Kunming. We also note that none of the three eastern cities, namely Shenzhen, Fuzhou and Xiamen, is cointegrated with each other or with the remaining central and western cities. This might indicate that in the long-run, the housing market conditions in eastern cities still significantly differ from the markets of the remaining cities, although much effort has been made to promote the integration process of Pan-PRD's economy.

TABLE 5.4 Pairwise convergence results with pre-specified coefficients (1,-1)

	Shenzhen	Fuzhou	Xiamen	Nanchang	Changsha	Nanning	Guiyang	Kunming	Chengdu
Shenzhen		-0.919	-1.838	-1.071	-0.875	-0.667	-1.077	0.563	1.040
Fuzhou	-0.642		-1.060	-0.796	-0.940	-1.009	-1.448	0.146	1.044
Xiamen	-2.109	-0.927		-0.840	-1.047	-0.880	-1.510	0.803	1.109
Nanchang	-0.853	-1.843	-0.733		-0.546	-0.612	-2.994**	0.206	0.588
Changsha	-1.144	-1.607	-1.529	-2.729		-0.648	-0.465	0.348	0.614
Nanning	-0.480	-2.144	-1.132	-2.090	-0.594		-0.577	0.174	0.392
Guiyang	-0.882	-2.170	-1.468	-3.075*	-2.900*	-2.451		-0.009	-0.056
Kunming	-0.480	-3.383*	-1.464	-3.550**	-4.012**	-3.770**	-2.949*		-1.154
Chengdu	0.509	-4.018**	-0.491	-1.810	-2.973*	-3.443*	-1.542	-2.124	

Notes: The null hypothesis of no convergence is tested based on the residual from $p_{it} - p_{jt}$. The results of ADF test (Dickey and Fuller 1979) are reported in the lower triangle, whereas the upper triangle shows the results of DF-GLS tests (Elliott et al. 1996). In the unit root test process, the constant is included, and the lag length (not reported) is automatically selected by the Bayesian Information Criterion (BIC). * and ** indicate the 5% and 1% significance level, respectively.

After examining the cointegration, we proceed to test the more restricted long-run convergence properties. The pairwise cointegration results with pre-specified coefficients (1,-1) are shown in Table 5.4. The ADF test in the lower triangle suggests that among the four cointegration city pairs, there are three convergent pairs: Nanchang-Guiyang, Changsha-Guiyang and Changsha-Kunming. In addition, another seven city pairs, which are not cointegrated in the Engle-Granger cointegration test are found to be significantly converged. Considering that cointegration is a necessary condition of convergence, the convergent results of these seven pairs are a surprise. These contradictory results might be due to the low power of the ADF test in detecting the unit root. To verify the ADF test, we also perform the more efficient DF-GLS test, the results of which are reported in the upper triangle of Table 5.4. This time we find only one significantly convergent pair, the Nanchang-Guiyang pair, which is also cointegrated. Note that regardless which unit root test we used, city pairs that are convergent are rare. Therefore, it is reasonable to conclude that cointegration or convergence is unlikely to widely exist among the nine cities, which indicates a diverged interurban housing market in the Pan-PRD.

To check the robustness of the cointegration and convergence test results based on the two-step procedure, we also conduct two additional tests: the pairwise Johansen cointegration test and the two-step convergence test based on the momentum threshold autoregressive (MTAR) unit root test (Enders and Granger 1998) which can allow an asymmetric adjustment. The trace statistics of the Johansen procedure, computed based on a VAR(3) specification with unrestricted intercept and no trend in VAR, are reported in the lower triangle of Table A1. The null hypothesis is that the column city is not cointegrated with the row city. For those cointegrated city pairs, the upper triangle of Table A1 reports the results of the log-likelihood ratio (LR) test, which is used to test the cointegrating vector restriction (1,-1). Table A2 displays the pairwise cointegration results with pre-specified coefficients (1,-1) based on the MTAR unit root test. Both of these two powerful methods identify a similar cointegration and convergence pattern between cities with the two-step procedure does; the results are largely in line with the lower triangle results of Table 5.4. Thus, we are confident of the previous finding that only very few city pairs are found to be cointegrated or convergent. For the following analysis, we mainly rely on the results of the two-step procedure.

§ 5.6 House price diffusion pattern

§ 5.6.1 Spatial-temporal house price diffusion model

Previous analysis suggests that most of the city pairs do not hold a long-run equilibrium relationship; however, a few pairs do. When modelling the house price dynamics of a city, we should consider the interrelation with both the cointegrated cities and the non-cointegrated cities. In other words, we should consider the influence from the cities that can impose a long-term effect and the cities that only have a transitory effect. In addition, the spatial dimension should also be considered because it is likely that the effect imposed by nearby cities is stronger than the influence of distant cities. The spatial-temporal house price diffusion model adopted in this paper can fully capture the characteristics along both spatial and temporal dimensions. This model is a variant of spatial-temporal diffusion model proposed by Holly et al. (2011) (the Holly model), which has been applied to investigating the effects of language border on the diffusion of house prices in Belgian markets by Helgers and Buyst (2016). However, unlike the Holly model, we do not designate a ‘dominant’ city, which is assumed to have contemporaneous effects on non-dominant cities. The reason for abandoning the ‘dominant’ city from our model specification is that the general lack of pairwise long-run cointegration relationships between cities found previously suggests that there is no city that can be seen as the long-run forcing for other cities. In our model specification, the house price series in the system excluding p_{it} is split into two groups because of the existence of a ‘cointegration club’: one group (denoted by C) being cointegrated with p_{it} and the other (denoted by O) not. A first order error correction specification for p_{it} is given by

$$\Delta p_{it} = \phi_{i0}\left(p_{i,t-1} - \beta_i \bar{p}^{C}_{i,t-1}\right) + a_i + a_{i1}\Delta p_{i,t-1} + b_{i1}\Delta \bar{p}^{C}_{i,t-1} + c_{i1}\Delta \bar{p}^{O}_{i,t-1} + \epsilon_{it}, \quad (4)$$

where $\bar{p}^C_{i,t-1}$ and $\bar{p}^O_{i,t-1}$ are the spatially lagged variables, defined by

$$\begin{cases} \bar{p}^C_{i,t-1} = \sum w_{ij} p_{jt}, & \text{if } p_{jt} \text{ belongs to cointegrating group} \\ \bar{p}^O_{i,t-1} = \sum w_{ij} p_{jt}, & \text{if } p_{jt} \text{ belongs to non-cointegrating group} \end{cases}$$

The weight, $w_{ij} \geq 0$, which describes the spatial interaction between city i and j, can be constructed either based on a contiguity measure or certain distance measures. Here, the weight is simply calculated by a simple inverse distance function

$$w_{ij} = 1/d_{ij} \tag{5}$$

where d_{ij} is the straightforward distance between the CBDs of city i and city j. In accordance with tradition, the weights are arranged in a row-standardized spatial weight matrix W.

Because p_{it} is cointegrated with each member in the cointegrating group C, it is expected to be cointegrated with $\bar{p}^C_{i,t-1}$ as well. The cointegrating parameter β_i can be estimated in advance and treated as known in estimating the equation (4). Even if city i has no cointegrated counterpart, the model can also be conducted by simply setting the error correction coefficient (ϕ_{i0}) to zero.

§ 5.6.2 Generalized impulse response function (GIRF)

After obtaining the parameter estimates of model (4) by ordinary least squares (OLS), we can construct the spatial-temporal impulse response functions for simulating and forecasting purposes. We begin by writing the system of equations (4) in matrix form

$$\Delta \mathbf{p}_t = \mathbf{a} + \boldsymbol{\Pi}\mathbf{p}_{t-1} + \boldsymbol{\Gamma}\Delta\mathbf{p}_{t-1} + \epsilon_t \tag{6}$$

where $\boldsymbol{\Gamma} = \mathbf{A}_1 + \mathbf{B}_1 + \mathbf{C}_1$, $\mathbf{p}_t = (p_{1t}, p_{2t}, \cdots, p_{nt})'$, $\mathbf{a} = (a_1, a_2, \cdots, a_n)'$, $\epsilon_t = (\epsilon_{1t}, \epsilon_{2t}, \cdots, \epsilon_{nt})'$,

$$\boldsymbol{\Pi} = \begin{bmatrix} \phi_{10} & 0 & \cdots & 0 & 0 \\ 0 & \phi_{20} & \cdots & 0 & 0 \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ 0 & 0 & \cdots & \phi_{n-1,0} & 0 \\ 0 & 0 & \cdots & 0 & \phi_{n0} \end{bmatrix} - \begin{bmatrix} \phi_{10}\beta_1 \mathbf{w}'_{1,C} \\ \phi_{20}\beta_2 \mathbf{w}'_{2,C} \\ \vdots \\ \phi_{n-1,0}\beta_{n-1} \mathbf{w}'_{n-1,C} \\ \phi_{n0}\beta_n \mathbf{w}'_{n,C} \end{bmatrix},$$

$$\mathbf{A}_1 = \begin{bmatrix} a_{11} & 0 & \cdots & 0 & 0 \\ 0 & a_{21} & \cdots & 0 & 0 \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ 0 & 0 & \cdots & a_{n-1,1} & 0 \\ 0 & 0 & \cdots & 0 & a_{n1} \end{bmatrix}, \mathbf{B}_1 = \begin{bmatrix} b_{11} \mathbf{w}'_{1,C} \\ b_{21} \mathbf{w}'_{2,C} \\ \vdots \\ b_{n-1,1} \mathbf{w}'_{n-1,C} \\ b_{n1} \mathbf{w}'_{n,C} \end{bmatrix}, \quad \text{and} \quad \mathbf{C}_1 = \begin{bmatrix} c_{11} \mathbf{w}'_{1,O} \\ c_{21} \mathbf{w}'_{2,O} \\ \vdots \\ c_{n-1,1} \mathbf{w}'_{n-1,O} \\ c_{n1} \mathbf{w}'_{n,O} \end{bmatrix}$$

where $\mathbf{w}'_{i,C}$ and $\mathbf{w}'_{i,O}$ represent the i_{th} row of spatial weight matrixes connecting with the cointegration group and the non-cointegration group, respectively. Matrix $\mathbf{A}_1$ indicates their own short-run influence, and matrixes $\mathbf{B}_1$ and $\mathbf{C}_1$ represent the short-run impacts of the cities from the cointegration group and the non-cointegration group, respectively.

The equation (6) can be rewritten as a form of vector autoregression (VAR)

$$\mathbf{p}_t = \mathbf{a} + \mathbf{\Phi}_1 \mathbf{p}_{t-1} + \mathbf{\Phi}_2 \mathbf{p}_{t-2} + \epsilon_t \tag{7}$$

where $\mathbf{\Phi}_1 = \mathbf{I}_n + \mathbf{\Pi} + \mathbf{\Gamma}$ and $\mathbf{\Phi}_2 = -\mathbf{\Gamma}$. The VAR model (7) can then be used for impulse response analysis. Suppose that the shock, ϵ_{it}, which will propagate to other cities, is characterised by the variance-covariance matrix

$$\Sigma = \begin{bmatrix} \sigma_{11} & \sigma_{12} & \cdots & \sigma_{1,n-1} & \sigma_{1n} \\ \sigma_{21} & \sigma_{22} & \cdots & \sigma_{2,n-1} & \sigma_{2n} \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ \sigma_{n-1,1} & \sigma_{n-1,2} & \cdots & \sigma_{n-1,n-1} & \sigma_{n-1,n} \\ \sigma_{n1} & \sigma_{n2} & \cdots & \sigma_{n,n-1} & \sigma_{nn} \end{bmatrix}$$

where $\sigma_{ij} = E\left(\epsilon_{it}\epsilon_{jt}\right)$, which can be consistently estimated from the OLS residuals $\hat{\epsilon}_{it}$ of the individual regressions, namely by $\hat{\sigma}_{ij} = T^{-1}\sum_{t=1}^{T}\hat{\epsilon}_{it}\hat{\epsilon}_{jt}$ and $\hat{\sigma}_{ii} = T^{-1}\sum_{t=1}^{T}\hat{\epsilon}_{it}^2$. To allow for possible contemporaneous correlation across cities, we consider the generalized impulse response function (GIRF) advanced in Pesaran and Shin (1998). The impulse response of a unit (one standard error) shock to house price in a city on the remaining cities at a horizon h periods ahead will be provided by

$$\begin{aligned} &\mathbf{g}_i\left(h\right) = E\left(\mathbf{p}_{t+h}|\epsilon_{it} = \sqrt{\sigma_{ii}}, \widetilde{S}_{t-1}\right) - E\left(\mathbf{p}_{t+h}|\widetilde{S}_{t-1}\right) = \frac{\mathbf{\Psi}_h \mathbf{\Sigma} \mathbf{e}_i}{\sqrt{\sigma_{ii}}} \\ &\text{for} \quad i = 1, \cdots, n; h = 0, 1, \cdots, H \end{aligned} \tag{8}$$

where $\widetilde{S}_{t-1}$ is the information set at time $t-1$ and $\mathbf{e}_i$ is an $n \times 1$ vector of zeros with the exceptions of its i_{th} element, which is unity, and

$$\mathbf{\Psi}_h = \mathbf{\Phi}_1 \mathbf{\Psi}_{h-1} + \mathbf{\Phi}_2 \mathbf{\Psi}_{h-2}, \tag{9}$$

with $\mathbf{\Psi}_0 = \mathbf{I}_n$ and $\mathbf{\Psi}_h = \mathbf{0}$ for $h < 0$.

§ 5.6.3 Empirical results

According to the previous cointegration test, four cities, Nanchang, Changsha, Guiyang and Kunming, are cointegrated with at least one of the other cities. For these four cities, we should include the error correction term $\left(p_{i,t-1} - \beta_i \bar{p}_{i,t-1}^C\right)$ in their house price dynamic specifications. In contrast, for the remaining cities, the error correction term and the term $\Delta\bar{p}_{i,t-1}^C$ in equation (4) can be eliminated.

The cointegration results between the four cities and the spatial lag of their cointegrated counterparts are shown in Table 5.5. As expected, the τ and z statistics provide significant evidence of cointegration for Nanchang, Changsha and Kunming. In the case of Guiyang, we cannot reject the null hypothesis of no cointegration at the 5% significance level; however, the statistics are marginally significant at the 10% level. The third column of Table 5.5 reports the estimated long-run relationships (β_i). The estimated β_i are approximately distributed around the value of unity, except for Kunming.

TABLE 5.5 Cointegration between the four cointegrating cities and spatial lag of their cointegrated counterparts

	τ statistic	z statistic	β
Nanchang	-3.561*	-25.128*	0.843441
Changsha	-4.541**	-34.150**	1.287945
Guiyang	-3.089	-18.209	0.871291
Kunming	-4.069**	-20.546*	0.602230

Notes: The test is based on one equation regression in which we take column cities as dependent variables and the spatial lags of their cointegrated counterparts as independent variables. The first two columns report the τ statistic and z statistic of ADF test with the null hypothesis of no cointegration. The third column reports the estimation of β. Note that the t-ratio for β is invalid in this case. * and ** indicate the 5% and 1% significance level, respectively.

With β_i being determined, the spatial-temporal house price dynamic model for each city can be estimated using ordinary least squares (OLS). The estimation results are summarized in Table 5.6 where the lag-orders are set to 2. These models perform reasonably well because the Breusch-Godfery test suggests no serial correlation in each regression's residuals at least at the 5% significance level.

The error correction terms, which appear in the model specification of four cities, are all significant at the 10% significance level, three of which are significant at the 5% level or above. That is, the four cities' short-run dynamics are influenced by the deviation from the long-run equilibrium relationship. The coefficient ϕ_{i0} indicates that the house price of Changsha responds to the disequilibrium much more rapidly than that of the other three cities.

We now turn to the influence of short-term dynamics. Not surprisingly, the first-lag price changes are significantly positive in all equations. The second-order lagged price dynamics also play a role in the price equation for Fuzhou and Xiamen. Similarly, the lagged price changes from neighbouring cities (either from the cointegration group or the non-cointegration group) are also found to be statistically significant in all equations, except for Changsha and Kunming. This confirms the existence of cross-city spillover effects from the neighbouring cities, which is in accordance with the findings of Holly et al. (2011) for the UK market and Helgers and Buyst (2016) for Belgian housing markets.

TABLE 5.6 Estimation results of the spatial-temporal house price diffusion model

City	Constant	Error correction	Own lag effects		Lag effects of cointegration group		Lag effects of non-cointegrated group		Adjusted R^2	Serial correlation
		ϕ_{i0}	a_{i1}	a_{i2}	b_{i1}	b_{i2}	c_{i1}	c_{i2}		
Guangzhou	0.0007	–	0.4326***	0.0765	–	–	0.6540***	-0.3067*	0.514	5.452
	(0.001)		(0.098)	(0.108)			(0.182)	(0.161)		
Shenzhen	0.0011	–	0.6464***	-0.0038	–	–	0.9324**	-0.7141***	0.527	6.035
	(0.001)		(0.106)	(0.106)			(0.197)	(0.200)		
Fuzhou	0.0005	–	0.3074***	0.2321*	–	–	0.5384***	-0.2231	0.548	0.588
	(0.001)		(0.100)	(0.097)			(0.140)	(0.141)		
Xiamen	0.0008	–	0.1673*	0.1624*	–	–	0.7223***	-0.1577	0.477	2.344
	(0.001)		(0.096)	(0.093)			(0.167)	(0.180)		
Nanchang	0.0548*	-0.0755*	0.3530***	0.0723	0.0737	-0.0436	0.3720**	-0.0430	0.425	4.708
	(0.028)	(0.040)	(0.104)	(0.108)	(0.098)	(0.096)	(0.171)	(0.174)		
Changsha	-0.2594***	-0.2006***	0.3654***	0.0726	0.1437	-0.0934	0.1763	0.1628	0.541	1.271
	(0.062)	(0.048)	(0.093)	(0.092)	(0.158)	(0.159)	(0.192)	(0.195)		
Nanning	0.0003	–	0.4869***	0.0244	–	–	0.3624**	0.0928	0.615	3.111
	(0.001)		(0.106)	(0.101)			(0.171)	(0.170)		
Guiyang	0.0414**	-0.0696**	0.4129***	-0.0258	0.4800***	-0.2064*	0.0976	-0.0413	0.467	7.778*
	(0.020)	(0.034)	(0.095)	(0.092)	(0.109)	(0.111)	(0.155)	(0.165)		
Kunming	0.1915***	-0.1092***	0.3423***	-0.0546	0.1323	-0.1476	0.3847	0.0062	0.258	0.855
	(0.059)	(0.034)	(0.102)	(0.098)	(0.109)	(0.106)	(0.263)	(0.232)		
Chengdu	-0.0002	–	0.2419**	0.0420	–	–	0.3315***	0.1753	0.550	1.934
	(0.000)		(0.099)	(0.098)			(0.120)	(0.124)		

Notes: The estimation results are based on equation (4) with lag-orders being set to 2. For the cities except Nanchang, Changsha, Guiyang and Kunming, we eliminate the error correction term and the term 'lag effects of cointegration group' because these cities do not have cointegrated counterparts. The "serial correlation" column reports the Breusch-Godfery serial correlation statistic test that distributes approximately as χ^2_3 under the null hypothesis of no residual serial correlation. *, ** and *** indicate the 10%, 5% and 1% significance level, respectively.

The regression results of spatial-temporal diffusion models presented in Table 5.6 depict a complicated dynamic system in which the historical house price changes of a city not only affect its own price changes but also influence the price changes in other cities directly or indirectly through their neighbouring cities, or through the long-run equilibrium. To intuitively illustrate the diffuse nature of house prices in a complicated system, we provide the generalized impulse response functions, which can trace the time profile of shocks both over time and space.

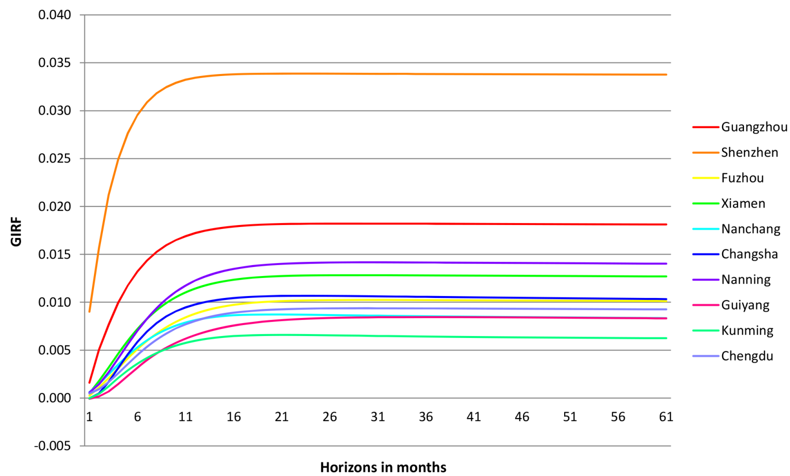


FIGURE 5.3 Generalized impulse responses of a positive unit shock (one standard error) to Shenzhen house prices

Figure 5.3 plots the generalized impulse responses of all the cities to a positive unit shock (one standard error) to the house prices of Shenzhen, one of the most developed cities in the Pan-PRD. The positive shock gradually diffuses to the remaining cities, significantly raising the house prices in the whole area (being confirmed by the bootstrap confidence interval in Figure B1 in the Appendix). However, the magnitude of the spillover effect differs across the region. Given the one standard error shock to Shenzhen, its own house prices soar approximately 3.5%, followed by Guangzhou, which rises by approximately 2%. Conversely, the increases of the other cities' house prices are approximately 1%. This indicates a diverged interurban housing market between developed and less-developed cities. For the sake of comparison, Figure 5.4 portrays the responses to a positive stand error shock to Changsha, a city in Central China and cointegrating with the other three cities. It is clear that the unit shock to Changsha generates relatively homogenous effects on all other cities' house prices (house prices increases are approximately between 0.8% and 1%), except for Shenzhen and Kunming. The effect on Shenzhen house prices is not significant, as indicated by the bootstrap error bounds shown in Figure B2 (refer to the Appendix). This information further supports our conclusion regarding the divergence of a few cities' housing markets, such as that of Shenzhen[12].

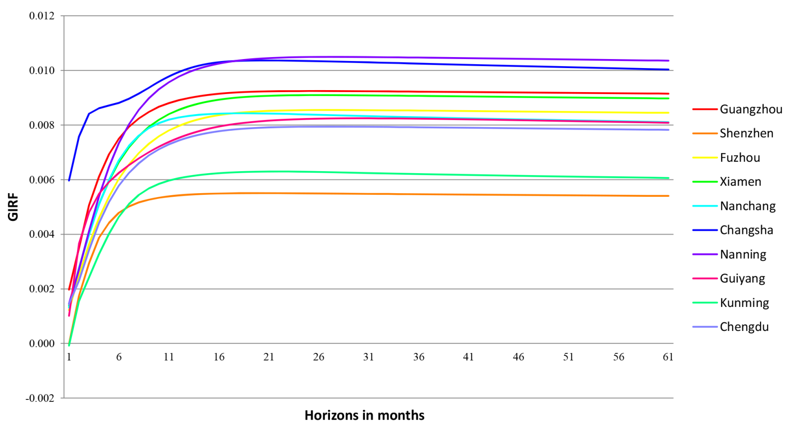


FIGURE 5.4 Generalized impulse responses of a positive unit shock (one standard error) to Changsha house prices

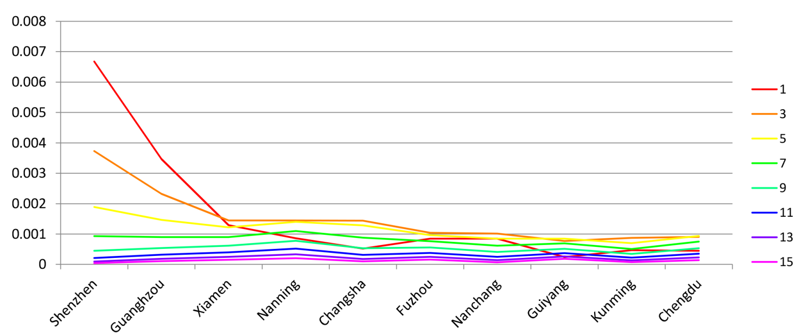


FIGURE 5.5 Generalized impulse responses of house price changes to one stand error shock to Shenzhen house prices

Figure 5.3 and Figure 5.4 also display a certain diffusion pattern in a manner that certain cities' response to shocks is more rapid than the others. To further examine the spatial-temporal diffusion pattern, Figure 5.5 depicts the impulse responses of house price changes to one standard error shock to Shenzhen house prices (the cities in the horizontal axis are ordered by distance). The first month after shock witnesses much higher house price increases in Shenzhen and its neighbouring cities than in cities far away. In the following few months, the house price changes of distant cities begin to

12 The impulse responses of the shock to other cities, which are not reported for space consideration, can lead to the similar finding that the overall interurban housing market is diverged. The responses are available upon request.

catch up, but remain slightly behind the neighbouring cities. Finally, house price changes in each city are nearly identical to each other after the seventh month. This behaviour clearly describes a diffusion pattern with the cities that are close to Shenzhen responding to shocks more rapidly and drastically.

§ 5.7 Conclusion and implications

Three aspects of the spatial interrelations of the 10 cities' housing markets in the Pan-PRD, namely spatial causality, convergence and diffusion, are carefully examined in this paper, based on the monthly house price indexes covering the period from June 2005 to May 2015. Among the 10 cities' housing markets, the Toda-Yamamoto Granger causality test reveals a complicated inter-market correlation pattern. It can be tentatively concluded that there is a causal flow from eastern-central China to the West considering that house prices of eastern-central cities are helpful in predicting house prices of western cities, but not vice versa.

In spite of the widely existing leading-lag interrelations, the Engle-Granger cointegration test provides very limited evidence for long-run cointegration among the cities. We find 4 cointegrated pairs of 36 city combinations. The evidence for convergence is rare too. Overall, the housing markets in the Pan-PRD are diverged, particularly between developed eastern and less developed western cities. The finding of divergence in the housing markets in Pan-PRD area contradicts most of the previous studies (e.g.,Wang et al. 2008; Huang et al. 2010b; Li and Li 2011) that support long-run cointegration of housing markets within a relatively homogeneous area. This suggests the possibility of 'club integration' and we indeed find a 'cointegration club' among the three cities in Central China.

In the short-run, the estimation results of the spatial-temporal diffusion model show that the house price change of a city can be influenced by its own lagged price changes, the spillovers from neighbouring cities, or the long-run forces from the cointegrated counterparts. Furthermore, the generalized impulse response functions (GIRF) confirm the divergence between developed and less developed housing markets because the shocks to Shenzhen can notably raise its own house prices but have limited influence on other cities' house price. However, a house price diffusion pattern can be conformed because the propagation of the shocks is approximately in accordance with the distance decay.

Similar to most of the studies on the Asian market, this paper is also limited by the short time-period, which is a notable issue when our analysis is concerned with long-run properties. This short time period of observation warns us that the results should be treated with caution. However, these results should have relevance to investors, policy makers and regulators. First, the leading-lag relationship among regional housing markets and the house price diffusion pattern could be useful for investors and portfolio managers to adjust their real estate portfolio accordingly. Second, a few implications can be drawn for policy makers and regulators. The lack of market convergence in the long-run could suggest that a local market-oriented housing policy will be more appropriate than a unified national policy. Indeed, this supposition has attracted the attention of policy makers. Recently, the central government announced a new policy to stimulate the housing market by reducing the down payment for second homes from 30% to 20%. An innovation of this policy is that it allows the local governments of four first-tier cities, Beijing, Shanghai, Shenzhen and Guangzhou, to make their own decisions according to the local market conditions[13]. Moreover, the results offer us a perspective on the degree of regional economic integration in the Pan-PRD. We observed that, during the following decade after the launch of Pan-PRD which aims to promote the integration of regional development, the developed eastern cities such as Shenzhen and Guangzhou, still appear to be deviated from the remaining cities, at least from the perspective of housing market integration. This suggests a need for regional policies that can facilitate the further decentralisation of economic activities, such as industrial policies.

References

Abbott, A., & Vita, G. D. (2013). Testing for long-run convergence across regional house prices in the UK: A pairwise approach. Applied Economics, 45(10), 1227-1238.

Alexander, C., & Barrow, M. (1994). Seasonality and cointegration of regional house prices in the UK. Urban Studies, 31(10), 1667-1689.

Ashworth, J., & Park, S. C. (1997). Modeling regional house prices in the UK. Scottish Journal of Political Economy, 44(3), 225-246.

Balcilar, M., Beyene, A., Gupta, R., & Seleteng, M. (2013). 'Ripple' effects in South African house prices. Urban Studies, 50(5), 876-894.

Berg, L. (2002). Prices on the second-hand market for Swedish family houses: Correlation, causation and determinants. International Journal of Housing Policy, 2(1), 1-24.

Blanco, F., Martín, V., & Vazquez, G. (2016). Regional house pice convergence in Spanish during the housing boom. Urban Studies, 53(4), 775-798.

Brady, R. R. (2014). The spatial diffusion of regional housing prices across U.S. states. Regional Science and Urban Economics, 46, 150-166.

Canarella, G., Miller, S., & Pollard, S. (2012). Unit roots and structural change: An

13 Document No [2015] 128 issued by Ministry of Finance of Peoples' Republic of China.

application to US house price indices. Urban Studies, 49(4), 757-776.
Chen, P.-F., Chien, M.-S., & Lee, C.-C. (2011). Dynamic modeling of regional house price diffusion in Taiwan. Journal of Housing Economics, 20(4), 315-332.
Clapp, J. M., Dolde, W., & Tirtiroglu, D. (1995). Imperfect information and investor inferences from housing price dynamics. Real Estate Economics, 23(3), 239-269.
Cook, S. (2003). The convergence of regional house prices in the UK. Urban Studies, 40(11), 2285-2294.
Dickey, D. A., & Fuller, W. A. (1979). Distribution of the estimators for autoregressive time series with a unit root. Journal of the American Statistical Association, 74(366), 427-431.
Drake, L. (1995). Testing for convergence between UK regional house prices. Regional Studies, 29(4), 357-366.
Elliott, G., Rothenberg, T. J., & Stock, J. H. (1996). Efficient tests for an autoregressive unit root. Econometrica, 64(4), 813-836.
Enders, W., & Granger, C. W. J. (1998). Unit-root tests and asymmetric adjustment with an example using the term structure of interest rates. Journal of Business & Economic Statistics, 16(3), 304-311.
Engle, R. F., & Granger, C. W. J. (1987). Co-integration and error correction: Representation, estimation, and testing. Econometrica, 55(2), 251-276.
Giussani, B., & Hadjimatheou, G. (1991). Modeling regional house prices in the United Kingdom. Papers in Regional Science, 70(2), 201-219.
Hanink, D. M., Cromley, R. G., & Ebenstein, A. Y. (2012). Spatial variation in the determinants of house prices and apartment rents in China. Journal of Real Estate Finance and Economics, 45(2), 347-363.
Helgers, R., & Buyst, E. (2016). Spatial and temporal diffusion of housing prices in the presence of a linguistic border: Evidence from Belgium. Spatial Economic Analysis, 11(1), 92-122.
Holly, S., Pesaran, M. H., & Yamagata, T. (2011). The spatial and temporal diffusion of house prices in the UK. Journal of Urban Economics, 69(1), 2-23.
Holmes, M. J., & Grimes, A. (2008). Is there long-run convergence among regional house prices in the UK. Urban Studies, 45(8), 1531-1544.
Holmes, M. J., Otero, J., & Panagiotidis, T. (2011). Investigating regional house price convergence in the United States: Evidence from a pair-wise approach. Economic Modelling, 28(6), 2369-2376.
Huang, F., Li, C., & Li, Y. (2010a). Ripple Effect of Housing Prices among Chinese Deputy Provincial Cities Based on an Alternative Approach. International Journal of Business Administration, 1(1), 19-24.
Huang, F., Zhou, Y., & Li, C. (2010b). Ripple effect of housing prices fluctuations among nine cities of China. Management Science and Engineering, 4(3), 41-54.
Kim, Y. S., & Rous, J. J. (2012). House price convergence: Evidence from US state and metropolitan area panels. Journal of Housing Economics, 21(2), 169-186.

Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., & Shin, Y. (1992). Testing the null hypothesis of stationary against the alternative of a unit root: How sure are we that economic series have a unit root. Journal of Econometrics, 54(1-3), 159-178.

Lean, H. H., & Smyth, R. (2013). Regional house prices and the ripple effect in Malaysia. Urban Studies, 50(5), 895-922.

Lee, C.-C., & Chien, M.-S. (2011). Empirical modeling of regional house prices and the ripple effect. Urban Studies, 48(10), 2029-2047.

Li, J., & Li, H. (2011). The spatial diffusion and convergence of urban housing price in the Pearl River Delta. Journal of Chongqing University (Social Science Edition), 17(6), 17-23.

Liao, W.-C., Zhao, D., Lim, L. P., & Wong, G. K. M. (2015). Foreign liquidity to real estate market: Ripple effect and housing price dynamics. Urban Studies, 52(1), 138-158.

Luo, Z. Q., Liu, C., & Picken, D. (2007). Housing price diffusion pattern of Australia's state capital cities. International Journal of Strategic Property Management, 11(4), 227-242.

MacDonald, R., & Taylor, M. P. (1993). Regional house prices in Britain: long-run relationships and short-run dynamics. Scottish Journal of Political Economy, 40(1), 43-55.

MacKinnon, J. G. (1996). Numerical distribution functions for unit root and cointegration tests. Journal of Applied Econometrics, 11(6), 601-618.

Meen, G. (1996). Spatial aggregation, spatial dependence and predictability in the UK housing market. Housing Studies, 11(3), 345-372.

Meen, G. (1999). Regional house prices and the ripple effect: A new interpretation. Housing Studies, 14(6), 733-753.

Miles, W. (2015). Regional house price segmentation and convergence in the US: A new approach. Journal of Real Estate Finance and Economics, 50(1), 113-128.

Pesaran, H. H., & Shin, Y. (1998). Generalized impulse response analysis in linear multivariate models. Economics Letters, 58(1), 17-29.

Pollakowski, H. O., & Ray, T. S. (1997). Housing price diffusion patterns at different aggregation levels: An examination of housing market efficiency. Journal of Housing Research, 8(1), 107-124.

Stevenson, S. (2004). House price diffusion and inter-regional and cross-border house price dynamics. Journal of Property Research, 21(4), 301-320.

Su, C. W., & Chang, H. L. (2013). Is income converging in China? A flexible Fourier stationary test. Economics of Transition, 21(2), 341-356.

Toda, H. Y., & Yamamoto, T. (1995). Statistical inference in vector autoregressions with possibly integrated processes. Journal of Econometrics, 66, 225-250.

Wang, S., Yang, Z., & Liu, H. (2008). Empirical study on urban housing price interactions among regional marekts in China. Research on Financial and Economic Issues, 6, 122-129.

Wu, J., & Deng, Y. (2015). Intercity information diffusion and price discovery in housing markets: Evidence from Google Searches. The Journal of Real Estate Finance and Economics, 50(3), 289-306.

Wu, J., Deng, Y., & Liu, H. (2014). House price index construction in the nascent housing market: The case of China. The Journal of Real Estate Finance and Economics, 48(3), 522-545.

Yunus, N., & Swanson, P. E. (2013). A closer look at the U.S. housing market: Modelling relationships among regions. Real Estate Economics, 41(3), 542-568.

Appendices

Appendix A

TABLE A1 The results of pairwise Johansen cointegration test

	Shenzhen	Fuzhou	Xiamen	Nanchang	Changsha	Nanning	Guiyang	Kunming	Chengdu
Shenzhen									
Fuzhou	7.60								1.95
Xiamen	8.69	7.44							
Nanchang	7.25	11.11	9.05		7.19**		0.16	1.57	
Changsha	10.02	9.83	7.27	20.27**			4.45*	8.80**	
Nanning	10.74	12.66	9.01	12.77	10.26			2.05	
Guiyang	6.81	14.52	10.30	18.03*	18.22*	13.74		0.01	
Kunming	2.40	13.20	6.46	16.89*	25.04**	18.55*	17.07*		
Chengdu	6.75	16.61*	4.12	15.46	13.93	12.94	15.03	13.22	

Notes: The lower triangle cells report the trace statistic of the pairwise Johansen cointegration test under the null hypothesis $H_0 : r = 0$. The test is based on a VAR(3) specification, with unrestricted intercept and no trend in VAR. For the co-integrated city pairs, the upper triangle cells report the log-likelihood ratio (LR) test for the cointegrating vector restriction (1,-1). * and ** denote 5% and 1% significance level, respectively.

TABLE A2 Pairwise convergence with pre-specified coefficients (1,-1) base on the MTAR unit root test

	Shenzhen	Fuzhou	Xiamen	Nanchang	Changsha	Nanning	Guiyang	Kunming	Chengdu
Shenzhen									
Fuzhou	2.08(1)								
Xiamen	3.72(1)	0.70(2)							
Nanchang	1.54(1)	1.43(1)	0.48(1)						
Changsha	3.19(1)	3.93(2)	1.40(2)	4.87(1)					
Nanning	0.18(1)	2.76(1)	1.96(2)	2.15(1)	0.32(1)				
Guiyang	1.90(1)	2.38(1)	1.55(2)	4.73(1)	4.66(1)	3.03(1)			
Kunming	0.91(1)	5.17*(1)	2.17(1)	7.81**(1)	8.29**(1)	6.01*(1)	4.10(1)		
Chengdu	0.24(1)	4.34(1)	4.60(2)	1.67(1)	3.72(1)	5.96*(1)	1.20(1)	2.71(1)	

Notes: The null hypothesis of no convergence is tested based on the residual from $p_{it} - p_{jt}$. In all the models a constant is included, and the lag length is shown in the parentheses. The 5% and 1% critical values are 5.02 and 7.10, respectively. * and ** denote 5% and 1% significance level, respectively.

Appendix B: Bootstrap GIRF confidence intervals

The methods for computing the bootstrap confidence intervals of the generalized spatial-temporal impulse response functions are borrowed from Holly et al.(2011). The estimated model (7) is first used to generate B bootstrap samples. The bth bootstrap sample can be obtained by the following Data Generation Process (DGP)

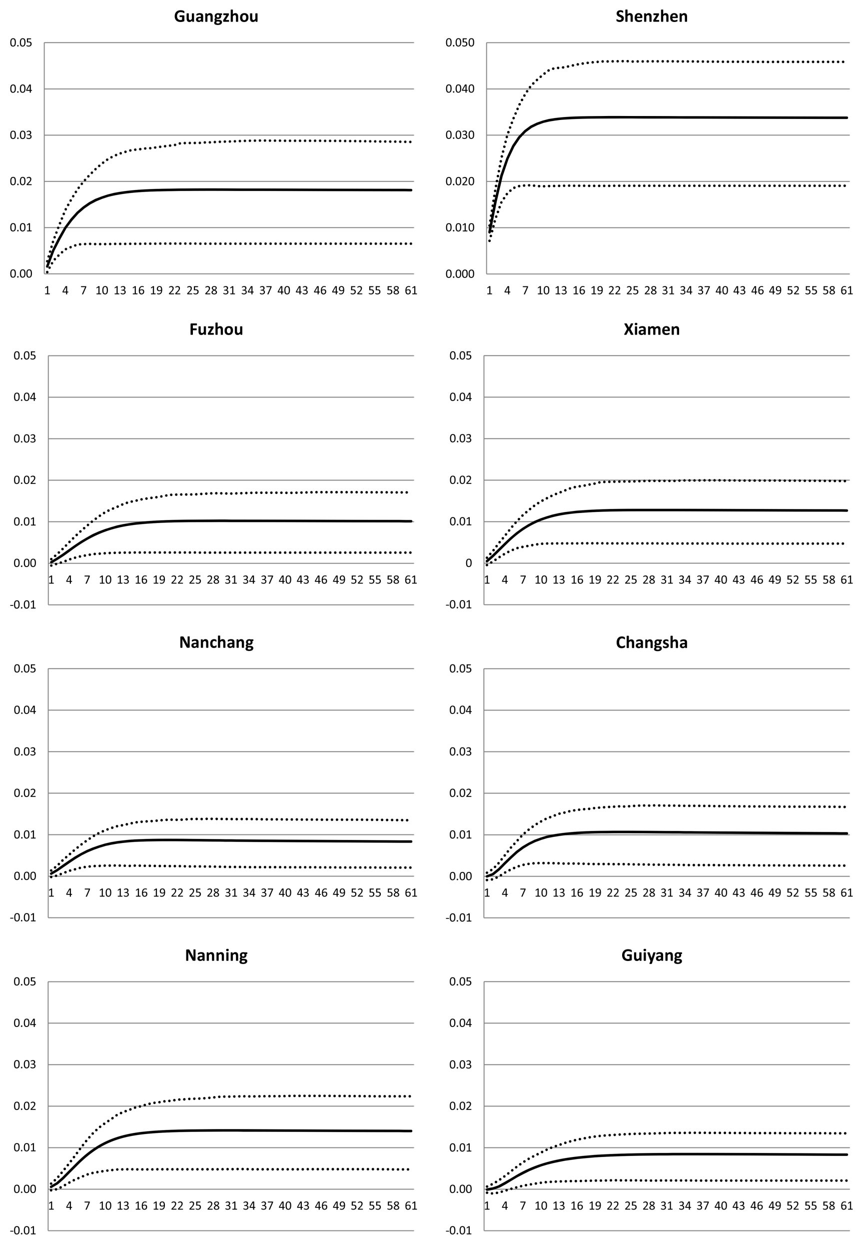
$$\mathbf{p}_t^{(b)} = \hat{\mathbf{a}} + \hat{\boldsymbol{\Phi}}_1 \mathbf{p}_{t-1}^{(b)} + \hat{\boldsymbol{\Phi}}_2 \mathbf{p}_{t-2}^{(b)} + \hat{\boldsymbol{\epsilon}}_t^{(b)}, \tag{B.1}$$

where $\hat{\boldsymbol{\epsilon}}_t^{(b)} = \hat{\boldsymbol{\Sigma}}^{1/2} \mathbf{v}_t^{*(b)}$. The elements of $\mathbf{v}_t^{*(b)}$ are recursively replaced by the values that are randomly drawn from the transformed residual matrix $\hat{\boldsymbol{\Sigma}}^{-1/2} (\hat{\boldsymbol{\epsilon}}_1, \hat{\boldsymbol{\epsilon}}_2, \ldots, \hat{\boldsymbol{\epsilon}}_t)$. Note that in equation (B.1), the first 2 observations are replaced by the original data.

When obtaining the bootstrap sample $\mathbf{p}_t^{(b)}$, we estimate the model (7) again and produce the b_{th} bootstrap GIRF

$$\mathbf{g}_i^{(b)}(h) = \frac{\hat{\boldsymbol{\Psi}}_h^{(b)} \hat{\boldsymbol{\Sigma}}^{(b)} \mathbf{e}_i}{\sqrt{\hat{\sigma}_{ii}^{(b)}}}, \quad \text{for} \quad i = 1, \cdots, n; h = 0, 1, \cdots, H. \tag{B.2}$$

The lower and upper bands of $100(1-\alpha)\%$ confidence interval are equivalent to the $\alpha/2$ and $1-\alpha/2$ quantiles of B $\mathbf{g}_i^{(b)}(h)$ for each i and h.


Guangzhou
0.05
0.04
0.03
0.02
0.01
0.00
1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
Shenzhen
0.050
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
Fuzhou
0.05
0.04
0.03
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
Xiamen
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
Nanchang
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
Changsha
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
Nanning
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61
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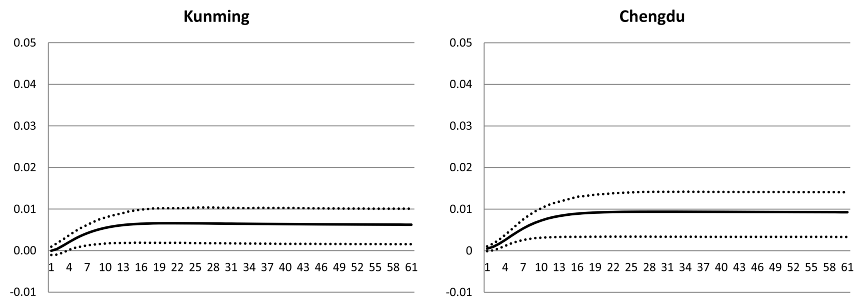
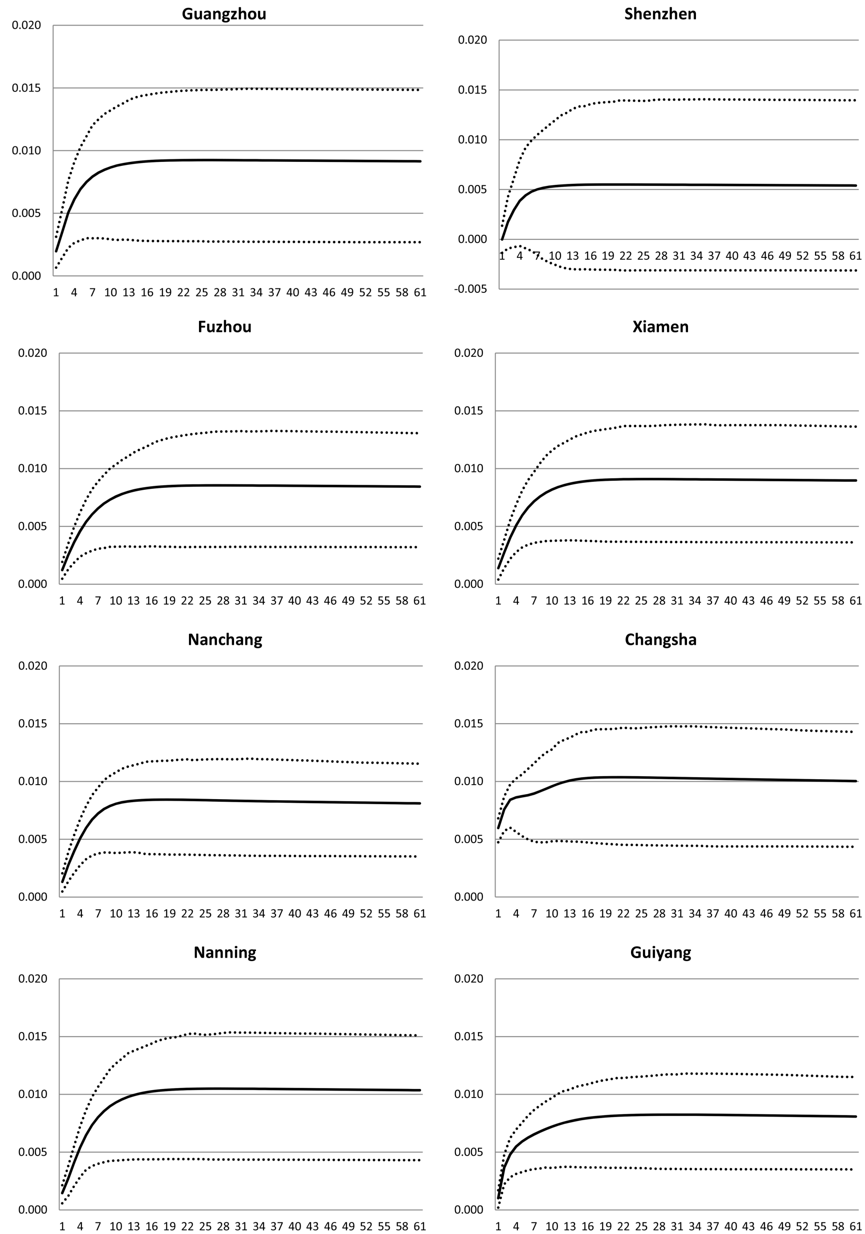


FIGURE B1 90% bootstrap error bounds for GIRF of a positive unit shock (one s.e.) to Shenzhen house prices (based on 1000 bootstrap samples)


Guangzhou
Shenzhen
Fuzhou
Xiamen
Nanchang
Changsha
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1 4 7 10 13 16 19 22 25 28 31 34 37 40 43 46 49 52 55 58 61

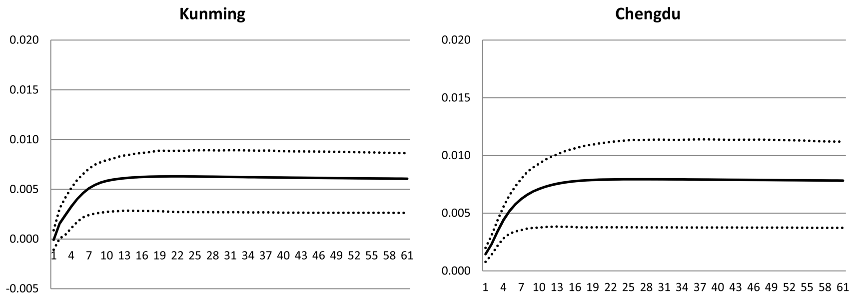

FIGURE B2 90% bootstrap error bounds for GIRF of a positive unit shock (one s.e.) to Changsha house prices (based on 1000 bootstrap samples)