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Research Article

A Linkage between New Commodity Home and Existing Home Prices in Chengdu, China

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Abstract: New home prices may mostly reflect long-run and investment factors. Existing home prices may mostly reflect current demand factors. This study implies a long-run equilibrium between these two types of home prices in Chengdu, a Chinese metropolis. The long-run elasticity of new commodity home prices with respect to existing home prices is 0.21. In the short run, the new home market leads the existing home market, which implies a comparatively healthy housing market. Despite this, two types of prices appear to contain a bubble at least during the period examined.
Keywords: house price, long run, new commodity home, existing home, cointegration, weak exogeneity.

INTRODUCTION

The housing market can be differentiated between new housing (first-hand home) market and existing housing (second-hand home) market. Both new home prices (NHP) and existing home prices (EHP) are influenced by the change in the number of households (a demand factor), economic growth, mortgage rate, inflation, land supply, current housing stock, future price expectations, etc. However, new home prices are most often impacted by price expectations while existing home prices are most often influenced by current stock and the current number of households (a current demand factor). Hence, new home prices may mostly be a reflection of long-run and even investment or speculative factors. Existing home prices may mostly be a reflection of current demand factors. An increase in new home prices could encourage higher existing home prices and vice versa. However, quick growth in new home and existing home prices represents different market settings; the former implies active housing investment, whereas the latter implies the real demand of current housing markets. This paper aims to test for the linkage between new home prices and existing home prices in Chengdu, China.

Chengdu is the capital of Sichuan Province in China. It is an essential central city in western China. In 2018, it had a land area of 14,335 square kilometers, accounting for 2.95% of Sichuan's total. It had a resident population of 16.33 million, accounting for 19.58% of Sichuan's total. The aggregate GDP reached RMB 1,534.3 billion (about 219.2 billion US dollars), accounting for 37.7% of Sichuan's total. Chengdu is located in the western part of the Sichuan Basin and the hinterland of the Chengdu Plain. It has a flat terrain and a river network. Chengdu holds developed agriculture. It has a subtropical monsoon humid climate. Chengdu is one of the most critical electronic information industry bases in the world, a quickly expanding business city in the world, and a benchmark city for investment in China's mainland.

METHODS

Cointegration implies a long-run equilibrium. To increase the test robustness, we examine cointegration using the Johansen trace and the Engle-Granger (or Phillips-Ouliaris) techniques (Engle R. F. *et al.*, 1987; Johansen S. *et al.*, 1990; Phillips P. C. B. *et al.*, 1990; Johansen S., 1991). Cheung-Lai and Reinsel-Ahn finite-sample corrections are used (Reinsel G. C. *et al.*, 1992; Cheung Y.-W. *et al.*, 1993). Unit root test applies ADF, PP and DF-ERS techniques (Dickey D. A. *et al.*, 1979; Phillips P. C. B. *et al.*, 1988; Elliott G. *et al.*, 1996). Also, the Zivot-Andrews test is used (Zivot E. *et al.*, 1992). We construct VAR or ECM models. Cointegrated variables suggest an ECM mechanism (Engle R. F. *et al.*, 1987). Variables with a unit root but no cointegration suggest VAR. We examine weak exogeneity in the cointegrated vector detected (Johansen S., 1992). We examine long-run and short-

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run elasticities in the ECM. We examine short-run elasticity in the VAR. We conduct Granger causality in either ECM or VAR (Granger C. W. J., 1969; Granger C. W. J., 1981).

Data

Monthly series cover the 2011-2015 period. House prices in Chengdu comprise existing home prices (*EHP*) and new commodity home prices (*NHP*). Prices are index changes as compared with the same month of last year. Data came from the National Bureau of Statistics of China (Nbsc, 2018).

Data were seasonally adjusted using the X-12 additive method. Logarithms were used. Table 1 presents a statistical description of the data. Figure 1 show that series have an intercept but may not contain a trend.

Table 1 Descriptive Statistics for the Raw Data

	EHP	*NHP*
Mean	99.77	101.48
Median	100.20	101.40
Max	105.30	109.80
Min	93.80	92.60
Std. Dev.	3.44	4.88
Skewness	-0.06	-0.02
Kurtosis	1.80	2.16
Jarque-Bera	3.65	1.75
p-value	0.16	0.42
Period	Jan 2011-Dec 2015	
Obs.	60	

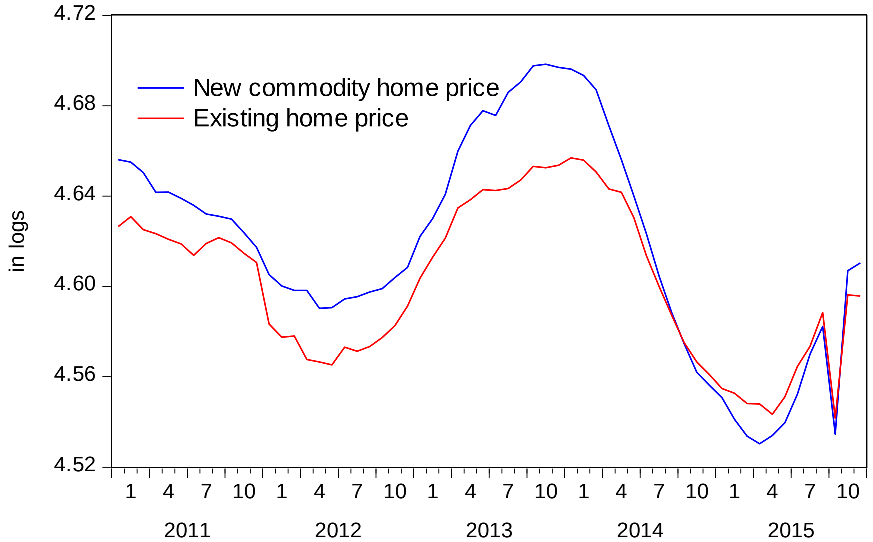

Figure 1 Monthly Change in Home Prices in Chengdu, China

Empirical Results

Unit Root

ADF tests suggested no unit root. PP tests suggested a unit root for two variables. DF-ERS tests implied more than one. Moreover, break-date tests suggested a unit root for two variables. Thus, this study treated these two variables as being I (1).

Table 1 the Unit Root Tests (ADF Tests)

Log variable	*k*	Level	*k*	First difference
EHP	5	-3.44**	-	-
NHP	4	-3.93***	-	-

Notes: All tests encompass an intercept. The lag length *k* was decided using the *t*-tes (Ng S. *et al.*, 1995). ** And ***denote rejection of the null of a unit root at the levels of 5% and 1 %, respectively.

Table 2 the Unit Root Tests (PP Tests)

Log variable	*k*	Level	*k*	First difference
EHP	4	-1.67	5	-8.62***
NHP	5	-1.64	5	-7.89***

Notes: All tests encompass an intercept according to (Hamilton J. D., 1994). The lag *k* was decided using the Newey–West (NW) bandwidth technique (Newey W. K. *et al.*, 1987). ***denotes rejection of the null of a unit root at the 1% level.

Table 3 The Unit Root Tests (DF-ERS Tests)

Log Variable	*k*	Level	*k*	First difference
EHP	0	-1.27	4	-1.58
NHP	0	-1.04	3	-1.23

Notes: Truncation lags, *k,* were chosen using the modified Akaike information criterion (MAIC). The MAIC is suggested to dominate all other criteria (Ng S. *et al.*, 2001). Test equations contained the intercept. Critical values used are in Table 1 (Elliott G. *et al.*, 1996).

Table 4 The Zivot-Andrews break-date test for EHP

		Coefficient	Standard Error	*t*-Statistic	p-value	T_{za}
Parameter	θ	0.01	0.01	1.38	0.17	
	β	0.00	0.00	1.55	0.13	
	γ	0.00	0.00	-2.95	0.01	
	α	0.38	0.15	2.61	0.01	-
k=9	1	-0.04	0.16	-0.27	0.79	-
	2	0.17	0.18	0.95	0.35	
	3	0.55	0.28	1.92	0.06	
	4	0.43	0.29	1.48	0.15	
	5	0.72	0.30	2.41	0.02	
	6	0.00	0.32	0.01	0.99	
	7	0.55	0.33	1.65	0.11	
	8	0.08	0.31	0.26	0.79	
	9	0.57	0.31	1.82	0.08	
	Constant	2.82	0.68	4.15	0.00	
	R-squared	0.94	Mean dependent var	4.60		
	Adjusted R-squared	0.92	S.D. dependent var	0.04		
	S.E. of regression	0.01	Akaike info criterion	-6.06		
	Sum squared resid	0.00	Schwarz criterion	-5.52		
	Log-likelihood	165.47	Hannan-Quinn criteria.	-5.86		
	F-statistic	43.56	Durbin-Watson stat	2.23		

Notes: Variable was in logarithmic values. Test equations included both a linear trend and a constant. The lagged length *k* (between 2 and 10) was selected using a general-to-specific recursive method. Thus, given lagged terms of variable, $x_{(t-k)}$, *t*-statistic on $x_{(t-k)} \geq 1.80$ but the term $x_{(t-(k+1))}$ is statistically insignificant. *k* was selected backward beginning from a maximum value of 10. This method is data-dependent. The trimming fraction is 0.29. The critical values for a sample of 71 were −6.25, −5.68, and −5.38 at 1%, 5%, and 10% levels, respectively (Zivot E. *et al.*, 1992). T_{za} is the possible break date selected.

Table 5 The Zivot-Andrews break-date test for NHP

		Coefficient	Standard Error	*t*-Statistic	p-value	T_{za}
Parameter	θ	0.01	0.01	1.00	0.32	
	β	0.00	0.00	0.16	0.88	
	γ	0.00	0.00	-1.67	0.10	
	α	0.72	0.09	8.23	0.00	-
k=4	1	-0.26	0.14	-1.93	0.06	-
	2	-0.05	0.20	-0.27	0.79	
	3	0.64	0.40	1.61	0.11	
	4	1.07	0.39	2.73	0.01	
	Constant	1.29	0.41	3.19	0.00	
	R-squared	0.95	Mean dependent var	4.62		
	Adjusted R-squared	0.94	S.D. dependent var	0.05		
	S.E. of regression	0.01	Akaike info criterion	-5.93		
	Sum squared resid	0.01	Schwarz criterion	-5.60		
	Log-likelihood	171.99	Hannan-Quinn criteria.	-5.80		
	F-statistic	116.72	Durbin-Watson stat	2.11		

Notes: The same as those in Table 4.

Cointegration

Both Engle-Granger and Johansen trace tests suggested cointegration. Notably, allowing for small sample, the Johansen test suggested a cointegrating vector. Hence, *NHP* and *EHP* are cointegrated. The normalized cointegration vector is

$$New\ home\ price = \underset{(-0.78)}{0.21} Existing\ home\ price + \underset{(-3.04)}{3.68}$$

The adjustment coefficient for *EHP* is -0.10. The adjustment coefficient for *NHP* is -0.16. Hence, both need to be adjusted downwards, which implied a price bubble.

Table 6 Engle-Granger Tests

Dependent variable	Z_α-statistic	p-value
NHP	−19.67	0.04
EHP	−22.11	0.02

Notes: Variables were in logarithms and first differences. Tests contained an intercept. Lags were chosen based on a t-statistic. p-values are provided in (Mackinnon J. G., 1996).

Table 7 Johansen cointegration trace tests

r	*k*	Eigenvalue	Trace	O-L*	C&L**	Reinsel-Ahn***
0	4	0.32	29.31	20.26	23.05	23.45
≤1		0.13	7.72	9.16	10.43	6.17

Notes: *r* is the null hypothesis of the cointegration rank of at most *r*. Models I, II, III, IV, and V are proposed for the trace statistic (Johansen S., 1991, 1995). We chose Model II (Hendry D. F. *et al.*, 2001). *5% Osterwald-Lenum asymptotical critical values (Osterwald-Lenum M., 1992). **5% Cheung-Lai finite-sample critical values (Cheung Y.-W. *et al.*, 1993). ***Reinsel-Ahn finite-sample trace corrections (Reinsel G. C. *et al.*, 1992). The lag length *k* was selected by reducing the Akaike information criterion (AIC) to the extent possible.

Weak exogeneity

For α_{11}=0, LR=9.66 with a p-value of 0.00, which rejected the weak exogeneity of *NHP* at the 1% level. For α_{21}=0, LR=4.09 with a p-value of 0.04, which implies we can accept the weak exogeneity of variable *EHP* at the 1% level.

Estimation of ECM

Having the cointegrating vector detected built into the first-differenced VAR, we estimated an ECM (Table 8). Regarding the short-run effect of *NHP* on *EHP*, the estimate on the third term is nearly significant (*t* statistic = 1.58) at the 10% level. ECM estimates would be used to test for Granger causality.

Table 8 ECM Estimates

		Estimate	*t*-statistic	Estimate	*t*-statistic
Error-correction term$_{t-1}$		-0.16	-3.75	-0.10	-2.31
	Lagged term	*NHP*		*EHP*	
NHP	t − 1	0.21	0.65	0.17	0.57
	t − 2	-0.07	-0.20	-0.13	-0.40
	t − 3	0.67	1.50	0.68	1.58
	t − 4	0.97	2.28	0.49	1.21
EHP	t − 1	-0.56	-1.60	-0.49	-1.47
	t − 2	.021	0.62	0.23	0.68
	t − 3	-0.49	-1.36	-0.13	-0.53
	t − 4	-0.02	-0.05	-0.19	-0.52
Constant	-3.68	-3.04			
R-squared	0.49				
Adj. *R*-squared	0.41				
F-statistic	5.60				
Akaike AIC	-5.95				

Granger causality

By excluding lagged *EHP* variables, χ^2 is 5.37 with a p-value of 0.25, which suggests no Granger causality from existing home prices to new home prices at the 10% level. By excluding lagged *New home price* variables, χ^2 is 8.25 with a p-value of 0.08, which suggests a Granger causality from new home prices to existing home prices at the 10% level.

Concluding Remarks

This study suggests a cointegration relationship between new commodity home prices and existing home prices in Chengdu. Hence, in the long run, these two types of prices move together. Since the existing home price is weakly exogenous, it has impacted the new home price in the long run. The long-run elasticity of new commodity home prices concerning existing home prices is 0.21.

The adjustment coefficients for the cointegrating vector are negative. The study suggests that both new commodity home prices and existing home prices in Chengdu may contain a price bubble.

In the short run, new home prices Granger caused existing home prices. Based on ECM estimates, the short-run elasticity of existing home prices with respect to new home prices is 0.68.

In all, in the long run, the existing home market impacts the new home market. In the short run, the new home market imposes an effect on the existing home market. However, both markets appear to contain a price bubble during the period examined.

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