Housing Returns, Inflation and Economic Growth:
Causality with structural breaks

2008-10-12

T.T.Binh, Nguyen
Department of Accounting, Chaoyang University of Technology, 41349, Taiwan, Taiwan

Kuan-Min Wang
Department of Finance, Overseas Chinese Institute of Technology, 100 Chiao Kwang Road, Taichung, 40721, Taiwan

Abstract
This paper investigates the housing-macroeconomic nexus in Taiwan with endogenous structural breaks during 1993~2006. GDP and CPI are taken into consideration for examining the inflation hedging ability of Taiwan’s housing returns and the contribution of housing market to economic growth. The empirical results show that the growth of GDP actually affects inflation, but it does not cause the growth in housing returns. Particularly, when taking the time trend into account, it is found that the effect of inflation on housing return is negative and the effect of housing return on inflation is positive. These evidences display the inflation hedging unavailability of Taiwan’s housing during studying period and the opportunistic characteristic of investors. Besides, the growth of housing market is not beneficial for economic growth in the long-run, yet it leads to higher inflation in the short-run.

Keywords: Housing Returns; Inflation Hedge; Economic Growth; Structural Breaks; Taiwan

JEL classification: C22, L85, P44
1. Introduction

Asset price fluctuations are generally considered to have impact on the real economy. Particularly, there are presumptions that the asset price movements are mirrored in the profile of economic activity, and that the duration and magnitude of the increase in asset prices, according to Kindleberger (2000), matter because they raise the vulnerability of the financial positions of households and firms to shocks. It is worth noting that real estate is regarded as a particular type of asset, similar to other assets, real estate prices are determined by many factors such as the expected service stream or expected future cash flow and the required rate of return as discount factor.

The role of real estate market in macroeconomy has attracted much attention among researchers and policymakers in recent years, yet historically, none of them address the causal and long-run relationship with endogenous structural breaks between real estate prices and macroeconomic variables. Structural breaks are, of course, important in investigating a long-run relationship between variables. In Perron (1989)’s opinion, most macroeconomic time series are best construed as stationary fluctuation around a deterministic trend function if allowance is made for the possibility of a shift in the intercept of the trend function. Clements and Hendry (1999) also state that failure to detect and account for parameter changes may lead to erroneous conclusions and poor forecasting performances. Accordingly, if there are breaks in the data, the results for a given replication may depend critically on whether and how often the breaks appear. Breaks can have a serious effect on stationarity of a series, the ignorance of possible shift, consequently, may yield misleading inferences.

The interaction between real estate prices and macroeconomic factors is evident in many studies. Apergis (2003) documents that housing loan rate is a variable with the highest explanatory power over the variation of real housing prices, followed by inflation and employment within the European Monetary Union. Besides, B.T. Ewing, J.E. Payne (2005) document, in the extent and the magnitude of the relationship between the real estate investment trusts (REITs) market and macroeconomic factors, that shocks to monetary policy, economic growth, and inflation all lead to lower than expected returns, while a shock to the default risk premium is associated with higher future returns. Accordingly, real estate markets are unable to avoid the influence of the economic hurricanes and big event occurred. The structural breaks of real estate prices, therefore, should present widely in many countries. Gerlach et al. (2006) recently discover the realty structural breaks of Japan, Malaysia, Hong Kong and Singapore. The break dates are about mid-to-late 1997, coinciding with the Asian
financial crisis.

According to Leung (2004), there is a relatively recent, growing recognition about the importance of the interactive nexus between and among housing markets and the macroeconomy. He also offers one question that how do the housing market and the macroeconomy intertwine? To answer this question with Taiwan’s case, we examine the causal relationship among Taiwan’s housing returns, CPI growth and real GDP per capita growth with Granger causality test. Since 1970, Taiwan real estate market has fluctuated along with economic growth and experienced three economic booms in 1973, 1979 and period from 1986 to 1989. The large fluctuation appears averagely in every seven years. Afterward, the real estate boom has vanished, which overturns the saying that “The boom of Taiwan real estate circulates in every seven years”. Taiwan’s real estate prices have dropped drastically since 1994 while the consumer price index and GDP per capita have been increasing stably.

Figure 1 illustrates Taiwan’s housing price, CPI and real GDP per capita as well as their growth rates which are known as the housing return, inflation rate and real GDP per capita growth rate over the period 1991Q3 to 2006Q2 when Sinyi housing index comes into use and reflects continuously the price dynamics. It is seen that the down gliding tendency of Taiwan’s housing prices appear obviously after 1994. It, even though seems to go up around 1996, continues with frustrated prices after 1997. In the meanwhile, Taiwan’ consumer price index and real GDP per capita show the up trends for the all sample period. Until 2003, housing prices, consumer price index and real GDP per capita just display the same up trend. This phenomenon shows that there potentially exits structural breaks in housing price series and that the long-run relationships between housing prices and Taiwan’ consumer price index as well as real GDP per capita are necessary to be further investigated.

To carry out the investigation, we devise a testing framework to examine if structural breaks influence the long-run relationship among housing prices (HPI), consumer price index (CPI) and real GDP per capita (GDP) over the period of 1991Q3 to 2006Q2. The relationships among the growth rates of these three variables are further examined with Granger causality test so as to verify the interaction among them. Our approach is divided into three basic steps: (1) carrying out the unit root tests as well as cointegration test with and without break for comparison, (2) exploring the long-run relationship among variables with cointegration test with a break, (3) finding the causal relationship among variables with a break.
The purpose of this paper is to provide evidence that the long-run relationship among Taiwan’s housing prices, GDP per capita and Consumer price index would be biased because the structural break is ignored. When the structural breaks are taken into account, the empirical results display the I(1) characteristic and the non-cointegration of these three variables. Therefore, three kinds of structural breaks are considered to create VAR model. The causal test shows that the growth of GDP actually affects inflation, but it does not cause the growth in housing return. In particular, when taking the time trend into account, it is found that the effect of inflation on housing return is negative and the effect of housing return on inflation is positive. These evidences display the inflation hedging unavailability of Taiwan’s housing during studying period and the opportunistic trait of investors. Besides, the growth of housing market is not beneficial to economy in the long-run, yet it leads to higher inflation in the short-run.

The remainder of this paper is organized as follows. Section 2 reviews some of the theoretical and empirical literature concerning the housing–macroeconomic nexus. Section 3 presents the data and summary statistics. Section 4 focuses on the methodologies and the main results on causal and long-run relationship with structural breaks among housing prices, growth of GDP per capita and Consumer price index. Finally, Section 5 concludes.

2. Literature Review

Starting with Veblen (1904) and Fisher (1933), the notion that asset prices might play a role in the macroeconomy generally is not new in economics. In the past decade, the notion seems to have become increasingly popular. The literature has furnished a great deal of studies concerning the realty–macroeconomic nexus and found extensive evidence that property price movements have a large impact on private consumption and the real economy.

Based on cross-country data for equity and real estate markets in most industrial countries, Higgins and Osler (1997) suggest that asset market bubbles during the late 1980s may have left the industrial world with an 'asset market hangover' in the early 1990s, in the form of sluggish asset markets and investment, which, of course, affect greatly the economic growth of industrial world. Focusing on GDP, Green (1997) shows that residential investment Granger causes GDP, but nor vice versa, while non-residential investment does not cause, but is caused by GDP. These results suggest that housing leads the business cycle in US.
Baffoe-Bonnie (1998) suggests that macroeconomic variables produce cycles in housing prices and houses sold. The housing market was found to be very sensitive to shocks in the employment growth and mortgage rate. He also argues that the economic variables have different impact on the dynamic behavior of housing prices and the number of houses sold in different regions at different time periods and that these economic aggregates alone cannot explain the fluctuations in real estate values and construction levels that occurred in some regions.

Furthermore, Iacoviello (2000) proposes that house prices can be embedded in a relatively simple macroeconometric model in a useful way, and that understanding their dynamics can shed some light over several macroeconomic episodes of the last quarter of century in Europe. According to Leung (2003), in Asia and North America, the economic growth persists result growth in real housing prices, even when population growth is zero and that the growth rate of a relative price index may converge to zero even when the housing prices in real terms display sustained growth.

The relationship between housing returns and inflation are discussed enthusiastically since one of the critical key pushing up the prices of real estate is its inflation hedge. Using conventional OLS models and cointegration and causality models to examine regional markets in the United Kingdom over 30 years, Stevenson (2000) finds strong evidence to support the hypothesis that housing and inflation are cointegrated and that housing leads inflation. Green concludes that policies to correct “over-investment” in housing to improve resource allocation in the long run context should be pursued with caution because doing so would have undesirable consequences in the short run. Recently, Kim (2004) investigates the relationships among housing price, consumer spending, and inflation in Korean economy. Particular attention is paid to the debate over house price bubbles, housing wealth effects on consumption, and the causality between house price and inflation. The test statistics suggest that the causality runs in both directions between house price increases and CPI inflation. In other words, not only does house price index provide useful information on inflation in the next month, inflation tends to lead to higher housing prices with the same lag.

The emergence of new financing methods in the past decade has made the discussion on relationship real estate and macroeconomy more excited. One of them is REITs of which the origins date back to the 1880s. However, REIT investment has become popular in academic literature for more than ten years. Being similar to
housing-macroeconomic nexus, many studies on REITs verify the relationship between REIT and macroeconomic variables. Chatrath and Liang (1998) find some evidence that REITs provide a long-run inflation hedge with Johansen (1988) test. However, the more standard residual-based cointegration techniques fail to provide similar evidence. Lu and So (2001) argue that the observed negative relationship between REITs returns and inflation is merely a proxy for the more fundamental relationship between REITs returns and other macroeconomic variables. Recently, Ewing and Payne (2005) document, in the extent and the magnitude of the relationship between the REITs market and macroeconomic factors, that shocks to monetary policy, economic growth, and inflation all lead to lower than expected returns, while a shock to the default risk premium is associated with higher future returns.

Despite of all these evidences of housing-macroeconomic nexus, there is surprisingly no evidence, to the best of our knowledge, taking structural breaks into account when investigating this long run relationship. The appearance of structural breaks in series, according to Zivot and Andrews (1992), can affect its stationary properties, and according to Gregory and Hansen (1996), can influence the long-run relationship of one set of variables.

3. Methodologies

To investigate both the role of housing market in macroeconomy and the effect of structural breaks, we employ development in econometrics on the properties of time series which has enabled researchers to investigate the relationship between integrated economic variables with ease and can provide precise estimates when structural break is taken into consideration. Firstly, two tests for unit root in a single time series, Dickey and Fuller (1981) and Phillips and Perron (1988), are used to test for stationarity. Test for unit root with a structural break developed by Zivot and Andrews (1992) is then used to compare to stationary results without consideration of structural break. Secondly, the cointegration between variables is further investigated when the stationary of series are confirmed. Two tests for cointegration, Johansen (1995a) and Gregory and Hansen (1996) or GH, are adopted. GH test with a Structural Break is applied to verify that the cointegration is biased if structural breaks are overlooked. If the cointegration exists, we utilize the error correction model to estimate the short-run adjustment. If the cointegration does not exist, we use the vector auto-regression model (VAR) for estimating. Thirdly, according to causal results of Granger (1986)
test, the causal direction and exogenous characteristic are determined. The three steps for testing unit root, cointegration and causality will be seen clearly below.

3.1 Tests for Stationarity

3.1.1 Conventional Linear Unit Root Tests

Among various conventional approaches for handling unit-root nonstationary, two approaches, Augmented Dickey-Fuller (ADF) and Phillips and Perron (PP) test, are selected for their popularity and representative. Figure 1 displays the effect of time trend on three variables, therefore, the testing mode that contains time trend is utilized, the Augmented Dickey-Fuller (ADF) tests are obtained as follows:

\[ \Delta y_t = \alpha + \phi y_{t-1} + \gamma t + \sum_{i=1}^{n-1} \beta_i \Delta y_{t-i} + \epsilon_t \]  

(1)

where, equation (1) is a test for random walk against stationary AR (1) and a deterministic trend. If \( H_0: \phi = 0 \) can not be rejected, the \( \{y_t\} \) series contains a unit root or \( \{y_t\} \) is non-stationary. The alternative hypothesis is \( H_1: -2 < \phi < 0. \)

Phillips and Perron have developed a more comprehensive theory of unit root non-stationarity. The test is similar to ADF test, but they incorporate an automatic correction to the DF procedure to allow for auto correlated residuals, which is known as PP test. The PP test first calculates the ADF statistics and modifies them to PP statistical values (Phillips-type test) under the circumstance that the error term is allowed to be of weak dependency and heterogeneous variance. The PP test is based on the statistic:

\[ \tilde{t}_a = t_a \left( \frac{\gamma_0}{f_0} \right)^{1/2} - \frac{T(f_0 - \gamma_0)(se(\hat{\alpha}))}{2f_0^{1/2}s} \]  

(2)

where \( \hat{\alpha} \) is the estimate, \( t_a \) is the t-ratio of \( \alpha \), \( se(\hat{\alpha}) \) is coefficient standard error, and \( s \) is the standard error of the test regression. In addition, \( \gamma_0 \) is a consistent estimate of the error variance in (1). The remaining term, \( f_0 \), is an estimator of the residual spectrum at frequency zero.

Since the estimation might be biased if the lag length and bandwidth are
pre-designated without rigorous determination, based on the “principle of parsimony”, our unit root tests utilize Akaike information criterion (AIC) and Bartlett kernel based criterion proposed by Newey and West (1994) for PP to determine optimal bandwidth.

3.1.2 Nonlinear Unit Root Test with a Structural Break

In performing unit root tests, special care must be taken if it is suspected that structural break has occurred. When there are structural breaks, Dickey-Fuller test statistics are biased towards the nonrejection of a unit root. Pierre Perron (1989) argues that structural breaks invalidate conventional unit root tests and develops a unit root model with an exogenous structural break. Afterwards, Zivot and Andrews (1992) upset the so-called exogenous break point and develop a unit root test (referred to as ZA test) with an endogenous structural break, which has been considered as a more suitable test for the stationarity of series. This study uses ZA test for the influence of structural break as well as of time trend on stationarity of level term. The equations ZA tests are described as the following forms:

\[ \Delta Y_t = \mu_t^D + \gamma_1^D t + \gamma_2^D DT_t^*, (\lambda) + \alpha^D Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \] (3)

\[ \Delta Y_t = \mu_t^C + \gamma_1^C t + \mu_2^C DU_t(\lambda) + \gamma_2^C DT_t^*(\lambda) + \alpha^C Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \] (4)

where \( DU_t(\lambda) \) equals 1 and \( DT_t^*(\lambda) = t - T\lambda \) if \( t > T\lambda \), and equals 0 otherwise.

\( \lambda = \frac{T_B}{T} \) and \( T_B \) represents a possible break point. Equation (3) allows for a break in the slope of trend of a series, while equation (4) allows for both breaks in the level and the slope of trend. The appropriate model and the optimal lag lengths of ZA tests utilize AIC.

3.2 Cointegration Tests

3.2.1 Johansen’s Cointegration Test

The concept of cointegration is interesting and has attracted a lot of attention in the literature. According to Engle and Granger (1987), a set of variables is defined as cointegrated if a linear combination of them is stationary. Many time series are
non-stationary but “move together” over time – that is, there exists some influences on the series, which imply that the two series are bound by some relationship in the long run. A cointegrating relationship may also be seen as a long-run or equilibrium phenomenon, since it is possible that cointegrating variables may deviate from their relationship in the short run, but their association would return in the long run.

The purpose of the cointegration test is to determine whether a group of non-stationary series are cointegrated or not. We implements VAR-based cointegration tests using the methodology developed in Johansen (1991, 1995a).

Consider a VAR of order $p$:

\[ y_t = A_1 y_{t-1} + \ldots + A_p y_{t-p} + Bx_t + \varepsilon_t \]  

where $y_t$ is a $k$-vector of non-stationary $I(1)$ variables, $x_t$ is a $d$-vector of deterministic variables, and $\varepsilon_t$ is a vector of innovations. We may rewrite this VAR as:

\[ \Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + Bx_t + \varepsilon_t \]  

where $\Pi = \sum_{i=1}^{p} A_i - I$ $\Gamma_i = -\sum_{j=i+1}^{p} A_j$.

Granger's representation theorem asserts that if the coefficient matrix $\Pi$ has reduced rank $\tau < k$, then there exist $k \times \tau$ matrices $\alpha$ and $\beta$ each with rank $\tau$ such that $\Pi = \alpha \beta'$ and $\beta' y_t$ is $I(0)$. $\tau$ is the number of cointegrating relations (the cointegrating rank) and each column of $\beta$ is the cointegrating vector. As explained below, the elements of $\alpha$ are known as the adjustment parameters in the VEC model. Johansen's method is to estimate the $\Pi$ matrix from an unrestricted VAR and to test whether we can reject the restrictions implied by the reduced rank of $\Pi$.

We consider the following five deterministic trend cases developed by Johansen (1995a, pp. 80-84):

1. The level data $y_t$ has no deterministic trends and the cointegrating equations do not have intercepts:

\[ H_2(\tau): \Pi y_{t-1} + Bx_t = \alpha \beta' y_{t-1} \]  

2. The level data $y_t$ has no deterministic trends and the cointegrating equations have intercepts:
3. The level data $y_t$ has linear trends but the cointegrating equations only have intercepts:

$$H_1^*(\tau): \Pi y_{t-1} + Bx_t = \alpha(\beta' y_{t-1} + \rho_0) \tag{8}$$

4. The level data $y_t$ and the cointegrating equations have linear trends:

$$H_1^*(\tau): \Pi y_{t-1} + Bx_t = \alpha(\beta' y_{t-1} + \rho_0 + \rho_t t) + \alpha_\perp \gamma_0 \tag{9}$$

5. The level data $y_t$ has quadratic trends and the cointegrating equations have linear trends:

$$H(\tau): \Pi y_{t-1} + Bx_t = \alpha(\beta' y_{t-1} + \rho_0 + \rho_t t) + \alpha_\perp (\gamma_0 + \gamma_t t) \tag{10}$$

The terms associated with $\alpha_\perp$ are the deterministic terms "outside" the cointegrating relations. When a deterministic term appears both inside and outside the cointegrating relation, the decomposition is not uniquely identified. Johansen (1995a) identifies the part that belongs inside the error correction term by orthogonally projecting the exogenous terms onto the $\alpha$ space so that $\alpha_\perp$ is the null space of $\alpha$ such that $\alpha' \alpha_\perp = 0$. We use a different identification method so that the error correction term has a sample mean of zero. More specifically, we identify the part inside the error correction term by regressing the cointegrating relations $\beta' y_t$ on a constant (and linear trend).

3.2.2 Nonlinear Cointegration Test with a Structural Break

Gregory and Hansen (1996) develop the test for cointegration which allow for the possibility of regime shifts. The possibility of unit roots being affected by structural change can also be the case with tests for cointegration. The GH test is a residual-based test, which test the null of no cointegration against the alternative of cointegration in the presence of a possible regime shift. The residuals are tested to determine if they are I(0) or I(1) processes. Rejection of the null implies an I(0) residual and, therefore, indicates the presence of cointegration. GH test can detect cointegrating relations when there is a break in the intercept and/or slope coefficient. Basically, the GH cointegration test incorporated with structural break is an extent of
Johansen’s linear model specification, and the test is regarded as a multivariate extension of the univariate ZA unit root test. The three models we use are described as follows:

\[(C)\] \[y_t = \mu_1 + \mu_2 D_t(\lambda) + \beta X_t + e_t\]  
\[(C/T)\] \[y_t = \mu_1 + \mu_2 D_t(\lambda) + \gamma t + \beta X_t + e_t\]  
\[(C/S)\] \[y_t = \mu_1 + \mu_2 D_t(\lambda) + \beta_1 X_t + \beta_2 X_t D_t(\lambda) + e_t\]  

where \(y_t\) denotes an univariate, and \(X_t\) represents variate vector. If \(t > T\lambda\) (\(1, \ldots, T\)), \(D_t(\lambda)\) equals 1, otherwise equals 0. \(\lambda = \frac{T_B}{T}\) and \(T_B\) denote a possible break point.

The test statistics for the ADF, and \(Z_a\) and \(Z_t\) of Perron (1989) are:

\[ADF^* = \inf_{\tau \in \tau} ADF(\tau), \quad Z^*_a = \inf_{\tau \in \tau} Z_a(\tau), \quad Z^*_t = \inf_{\tau \in \tau} Z_t(\tau)\]

The residuals are then tested to determine whether they are I(0) processes. When the residuals are I(0) processes, indicating the existence of cointegration.

### 3.3 Granger Causality Test

If the variables are non-stationary and cointegrated, the adequate method to examine the causal relations is the Vector Error Correction Model (VECM) (Granger, 1988); otherwise a VAR model is used in the case of no cointegration found among the variables (Granger, 1969).

The standard **Granger Causality** test then examines if there exists bilateral or unilateral causality between variables. The GH cointegration test can verify if structural break affects the long-run relationship between variables. Hence, structural break should be taken into account when proceeding causal test. In order to obtain a general result and related comparison, we take account of four VAR models\(^1\). The first

---

\(^1\) Generally, when the cointegration between variables exists, it is necessary to additionally consider the error correction term of previous period- viz. long-run cointegration vector- for analysis. Because the empirical results of this study support the non-cointegration between variables. We, therefore, use VAR model to test for causal effect.
VAR model that contains three series $X_t$, $Y_t$ and $Z_t$ without structural breaks is named model A:

\[
\Delta X_t = \alpha_1 + \sum_{i=1}^{n_1} \alpha_{11} (i) \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{12} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_1} \alpha_{13} (k) \Delta Z_{t-k} + \varepsilon_{1t} \tag{15}
\]

\[
\Delta Y_t = \alpha_2 + \sum_{i=1}^{n_2} \alpha_{21} (i) \Delta X_{t-i} + \sum_{j=1}^{m_2} \alpha_{22} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_2} \alpha_{23} (k) \Delta Z_{t-k} + \varepsilon_{2t} \tag{16}
\]

\[
\Delta Z_t = \alpha_3 + \sum_{i=1}^{n_3} \alpha_{31} (i) \Delta X_{t-i} + \sum_{j=1}^{m_3} \alpha_{32} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_3} \alpha_{33} (k) \Delta Z_{t-k} + \varepsilon_{3t} \tag{17}
\]

The second VAR model that contains the structural break of constant term is named model B:

\[
\Delta X_t = \alpha_1 + \sum_{i=1}^{n_1} \alpha_{11} (i) \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{12} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_1} \alpha_{13} (k) \Delta Z_{t-k} + \psi_1 D_t (\lambda) + \varepsilon_{1t} \tag{18}
\]

\[
\Delta Y_t = \alpha_2 + \sum_{i=1}^{n_2} \alpha_{21} (i) \Delta X_{t-i} + \sum_{j=1}^{m_2} \alpha_{22} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_2} \alpha_{23} (k) \Delta Z_{t-k} + \psi_2 D_t (\lambda) + \varepsilon_{2t} \tag{19}
\]

\[
\Delta Z_t = \alpha_3 + \sum_{i=1}^{n_3} \alpha_{31} (i) \Delta X_{t-i} + \sum_{j=1}^{m_3} \alpha_{32} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_3} \alpha_{33} (k) \Delta Z_{t-k} + \psi_3 D_t (\lambda) + \varepsilon_{3t} \tag{20}
\]

The third VAR model that contains the structural break of constant term and time trend is named model C:

\[
\Delta X_t = \alpha_1 + \sum_{i=1}^{n_1} \alpha_{11} (i) \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{12} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_1} \alpha_{13} (k) \Delta Z_{t-k} + \gamma_1 t + \psi_1 D_t (\lambda) + \varepsilon_{1t} \tag{21}
\]

\[
\Delta Y_t = \alpha_2 + \sum_{i=1}^{n_2} \alpha_{21} (i) \Delta X_{t-i} + \sum_{j=1}^{m_2} \alpha_{22} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_2} \alpha_{23} (k) \Delta Z_{t-k} + \gamma_2 t + \psi_2 D_t (\lambda) + \varepsilon_{2t} \tag{22}
\]

\[
\Delta Z_t = \alpha_3 + \sum_{i=1}^{n_3} \alpha_{31} (i) \Delta X_{t-i} + \sum_{j=1}^{m_3} \alpha_{32} (j) \Delta Y_{t-j} + \sum_{k=1}^{l_3} \alpha_{33} (k) \Delta Z_{t-k} + \gamma_3 t + \psi_3 D_t (\lambda) + \varepsilon_{3t} \tag{23}
\]

In equation (15)–(17), (18)–(20), (21)–(23), $\varepsilon_{xt}$, $\varepsilon_{yt}$ and $\varepsilon_{zt}$ are stationary random processes intended to capture other pertinent information not contained in lagged values of $X_t$, $Y_t$ and $Z_t$. The lag lengths, $n_i$, $m_j$ and $l_k$, are decided by AIC. $D_t (\lambda)$ is the dummy variable of structural breaks, the setting of dummy variables are displayed in equation (12), (13) and (14). According to equation (15) and (18), the series $Y_t$ fails to Granger cause $X_t$ if $\alpha_{12} (j) = 0$ $(j=1,2,3,..., m_1)$ and $Z_t$ fails to Granger cause $X_t$ if
\( \alpha_{13}(k) = 0 \) (\( k=1,2,3,\ldots, l_1 \)); according to equation (16) and (19) the series \( X_t \) fails to cause \( Y_t \) if \( \alpha_{21}(i) = 0 \) (\( i=1,2,3,\ldots, n_1 \)) and \( Z_t \) fails to Granger cause \( Y_t \) if \( \alpha_{32}(j) = 0 \) (\( j=1,2,3,\ldots, m_1 \)); according to equation (17) and (20), the series \( X_t \) fails to cause \( Y_t \) if \( \alpha_{31}(i) = 0 \) (\( i=1,2,3,\ldots, n_3 \)) and \( Y_t \) fails to Granger cause \( Z_t \) if \( \alpha_{23}(k) = 0 \) (\( k=1,2,3,\ldots, l_2 \)).

The final VAR model that considers two different regimes is named model D:

\[
\Delta X_t = [1 + \psi_1 D_t(\lambda)] [\alpha_1 + \sum_{i=1}^{n_1} \alpha_{11}(i) \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{12}(j) \Delta Y_{t-j} + \sum_{k=1}^{l_1} \alpha_{13}(k) \Delta Z_{t-k}] + \varepsilon_{Xt} \tag{24}
\]

\[
\Delta Y_t = [1 + \psi_2 D_t(\lambda)] [\alpha_2 + \sum_{i=1}^{n_1} \alpha_{21}(i) \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{22}(j) \Delta Y_{t-j} + \sum_{k=1}^{l_1} \alpha_{23}(k) \Delta Z_{t-k}] + \varepsilon_{Yt} \tag{25}
\]

\[
\Delta Z_t = [1 + \psi_3 D_t(\lambda)] [\alpha_3 + \sum_{i=1}^{n_1} \alpha_{31}(i) \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{32}(j) \Delta Y_{t-j} + \sum_{k=1}^{l_1} \alpha_{33}(k) \Delta Z_{t-k}] + \varepsilon_{Zt} \tag{26}
\]

\( D_t(\lambda) = 0 \) belongs to pre-break period. \( D_t(\lambda) = 1 \) belongs to post-break period. Because of the appearance of two regimes, the pre-break causal relationships in equation (24), (25), (26) are the same as the causalities in equation (15)~(23). According to equation (24), the post-break causal relationship displays that the series \( Y_t \) fails to Granger cause \( X_t \) if \( [1 + \psi_1] \alpha_{12}(j) = 0 \) (\( j=1,2,3,\ldots, m_1 \)) and \( Z_t \) fails to Granger cause \( X_t \) if \( [1 + \psi_1] \alpha_{13}(k) = 0 \) (\( k=1,2,3,\ldots, l_1 \)); According to equation (25),the series \( X_t \) fails to cause \( Y_t \) If \( [1 + \psi_2] \alpha_{21}(i) = 0 \) (\( i=1,2,3,\ldots, n_1 \)) and \( Z_t \) fails to Granger cause \( Y_t \) if \( [1 + \psi_2] \alpha_{23}(k) = 0 \) (\( k=1,2,3,\ldots, l_2 \)); According to equation (26), the series \( X_t \) fails to cause \( Y_t \) if \( [1 + \psi_3] \alpha_{31}(i) = 0 \) (\( i=1,2,3,\ldots, n_3 \)) and \( Y_t \) fails to Granger cause \( Z_t \) if \( [1 + \psi_3] \alpha_{32}(j) = 0 \) (\( j=1,2,3,\ldots, m_3 \)).

Granger causal test can further help determine the degree of exogenous characteristic as well as the causal direction.

4. Data Description and Empirical Evidence

Our analysis is based on the quarterly data of housing index (HPI), consumer price index (CPI) and real GDP per capita (GDP). The housing index is obtained from Taiwan Sinyi Realty Company, the consumer price index and real GDP per capita are obtained from Taiwan AREMOS database for the period from the third quarter of
Figure 3 displays the time trends of three variables in level term and first differential form. It is observed that HPI slides down during 1991-2002 and climbs up during 2002-2006. Whereas, CPI and real GDP per capita that are on the rise mostly presents the opposite direction to HPI in pre-2000 and the same rising direction to HPI in post-2000. The trend and volatility of these two variables and HPI are not stable not only in short run but also in long-run, which is changeable with times and produces different effects. Additionally, the first differential also means the growth rate. Therefore, based on its variation, the breadth of HPI growth is observed within -8% and 4% and its negative values are comparatively lager than positive values. However, the breadth of CPI growth is within -2% and 2% and the breadth of GDP growth is within -4% and 8%. The positive values of CPI and GDP are comparatively lager than positive values. According to figure 1, the three variables, more or less, present structural breaks. For example, the period of 994~1995, 1997~2000 and 2001~2003 connotes the existence of possible structural breaks which might affect the stationary as well as long-run relationship of variables. Hence, the disregard of these structural breaks will cause errors in estimation.

Table 1 reports the mean and standard deviation of growth. The mean value of HPI is negative, indicating that housing return is negative during studying period. Meanwhile, the mean value of CPI and GDP per capita are positive. Risk represented by standard deviation shows the higher risk of GDP per capita and less risk of CPI. The skewness values display left-skewed of three variables and the kurtosis value of HPI is higher than 3, exhibiting the form of leptokurtic, the kurtosis value of CPI, GDP exhibits the form of platykurtic

Starting with empirical process, we first look at the stationary of the series. The three variables are formed in natural logarithm while carrying out the test as well as the estimation for examining the relationship among them. Visual inspection suggests that HPI, CPI and real GDP per capita are nonstationary. These three series display

---

2 Sinyi housing index is established by Sinyi Realty Company that presently is the biggest real estate brokerage In Taiwan. Using the transaction database of existing homes in four biggest metropolises Taipei city, Taipei County, Taichung City and Kaohsiung City, its housing index is built up by the Hedonic Price Theory. Sinyi housing index came into effect in 1991, being the earliest index in Taiwan relative to two other indices including the north area housing index and Cathay housing index. Hence, the data that is provided by Sinyi housing index is not only the longest but also widely referenced as the representative benchmarking. It amply reflects the variational trend of the whole Taiwan’s housing prices. Additionally, the housing price index is adjusted quarterly based on the actual transactions of housing market.
structural breaks non-synchronously. ADF and PP are applied to test for stationarity on both the level and the first differential of the series. The results tabulated in Table 2 clearly suggest that none of the variables are stationary at level, that is, they are integrated of order 0. However, they are all characterized as integrated of order 1, that is, first differential will render these series stationary. We thus conclude that all variables considered in this paper are I(1) series. The summary statistics of ZA tests are presented in Table 3 that displays the non-stationarity of all series when a structural break is allowed. These results are consistent with those of ADF. The same order of integration between non-stationary series found from both linear and nonlinear tests enable us to go further with cointegration tests.

Testing for cointegration with Johansen approach, we diagnose models one by one until the model that cannot be rejected for the null. The empirical findings for the long-run relationship considering of a linear trend and a quadratic trend among GDP, CPI and HPI are summarized in Table 4. When rank $r = 0$ all Trace-test statistics of five models show the significant cointegration between two variables. When rank $r \leq 1$ the Trace-test statistics are significant in model 1 and model 5. When rank $r \leq 2$ all Trace-test statistics are not significant. The AIC criterion is adopted in selecting the appropriate lag lengths of 3 for our cointegration test. Based on the decision procedure described above, it is found that there significantly exists co-movement or long-run relationship among GDP, CPI and HPI.

In order to verify if structural break influences these long-run relationships, the GH Test that allows for one structural break is applied and its results are shown in Table 5. All coefficients of models are not significant, in other words, the existence of co-movement or long-run relationship among GDP, CPI and HPI are not clear. These finding presents the non-existence of long-run relationship between the two variables and support the suggestion of Clements and Hendry (1999) that failure to detect and account for structural break may lead to erroneous conclusions and poor forecasting performances.

It is obviously seen in figure 1 that the negative growth of CPI first appears from 1994 to 1995, then from 1997 when Asian financial crisis occurred until the mid of 2000. At that time, the stock market bubbles were evident worldwide, which seriously affected Taiwan’s economic growth. Beside, after experiencing SARS\(^3\) event in 2003, the annual economic growth decreased to 3.24%, 0.35% less than 3.59% of 2002. The appearance of structural breaks is found not only in inflation, economic growth, but also in housing price of Taiwan. Overlooking these structural breaks might lead to the

---

\(^3\) Severe Acute Respiratory Syndrome.
biased empirical results.

According to GH cointegration test, structural break evidently affects the relationship between variables. Therefore, it is necessary to consider the effect of structural break in testing the causality. The structural breaks detected by GH cointegration are used as the dummy variables for analysis. Except for equations (15)–(17) of model A, other related models all take structural break into account, equations (18)–(20) of model B allow for the structural break of constant, equations (21)–(23) of model C allow for the structural break of constant and time trend, equations (24)–(26) of model D allow for the structural break of constant and switch. The results of Granger’s Causality test are summarized in Table 6.

The causality tested with model A shows that returns of HPI positively and unilaterally affect the growth of CPI, the bilateral or feedback causality is discovered between the growth of GDP and CPI. The causality tested with model B indicates that the growth of GDP affect the growth of CPI positively and unilaterally, yet insignificantly. The causality tested with model C displays that causal relationship between the growth of CPI and returns of HPI is bilateral; the growth of CPI negatively affects the returns of HPI, and the returns of HPI have positive impact on the growth of CPI, the growth of GDP negatively and unilaterally affects the growth of CPI. Finally, the causality tested with model D presents the bilateral or feedback causal relationship between the growth of GDP and CPI in pre-break period and non-causal relationship between variables in post-break period.

In order to further ensure the appropriateness of structural breaks as well as the time trend in models, we utilize t test to examine each variable and Joint test to inspect if structural break or time trend should present in models. For model B, the results of t test displays the increase of conditional mean of housing returns due to the tempest of SARS in March 2003, the negative growth of inflation after April 1994, and the insignificant negative growth of economic growth after April 2002. However, the results of Joint test do not support the structural break of constant in model. For model C, the results of t test show that after SARS tempest, although the effect of time trend is negative (-0.001), the value of dummy variable coefficient of constant is higher (0.043). In sum, the conditional mean of housing returns increases, whereas, inflation negatively grows. Additionally, due to 1997 Asian financial crisis, Taiwan’s economy grows negatively, yet insignificantly. The results of Joint test support the structural break of constant as well as the time trend of variable. For D model, the t test shows that there was a significant drop of housing returns during the period of world stock market bubble in January 2001, and that inflation negatively grew in 1995, whereas, economic growth was insignificantly negative in March 2003. Lastly, the
results of Joint test support the existence of two regimes in model. The causal directions between variables of model 1 to model 4 are summarized in table 7.

All main empirical results are displayed in table 6 and table 7. Model 3 and 4 better present the causal relationship between inflation and economic growth, which has been approximately positive before 1995, implying that economic growth drives the increase of inflation. However, even though economic growth and inflation, as displayed in figure 1, both have had decreasing tendency after 1995, the variation of inflation is more significant. Therefore, economic growth of this period negatively affects inflation. The relationship between economic growth and housing returns with or without structural breaks is not the direct causality. The increase in housing returns just leads to higher inflation, implying that the investing or opportunistic element is in the majority. This phenomenon is helpless to promote economic growth in long-run as well as to hedge against inflation. Hence, the time trend within studying period affects the stationary of three variables and the long-run relationship between them. Although inflation decreases when considering time-varying process, we find the fact that Taiwan’s economy turns down, housing market is in recession, and housing returns are unable to hedge against inflation.

The empirical results of this study are obviously different from those of Green (1997), Chatrath and Liang (1998) Stevenson (2000), yet close to findings of Kim (2004) who investigates the relationships among housing price, consumer spending, and inflation in Korean economy. However, this study allows for the endogenous structural break of series to provide different information for analyzing the risk and uncertainty that Taiwan’s housing investment might face.

**Concluding remarks**

In many countries, asset price developments are considered to have contributed to macroeconomic stability. This applies not only to development on markets for financial assets, such as stock markets, but also to developments in real estate markets. Therefore, recognizing how real estate prices might influence one economy is crucial.

This paper makes two important points: 1. the long-run cointegration among CPI, HPI, and real GDP per capita does not exist when structural break is taken into consideration; 2. The causal test shows that the growth of GDP has effect on inflation, and that the causality between the growth of GDP and housing returns does not exist; 3. when taking the existence of time trend into account, the effect of inflation on housing returns presents a negative sign, and the effect of housing returns on inflation
presents a positive sign. We, thereby, conclude that Taiwan’s housing returns are not able to hedge against inflation because there are so many opportunists rather than demanders. Additionally, although housing growth is not helpful to economic growth in the long-run, it drives higher inflation in the short-run.

While many studies illuminates many possible responses of macroeconomy to housing prices, our empirical evidence provide more direct results that help policymakers to have a clearer look inside the economic position of Taiwan’s housing market. In fact, Kindleberger (1995) points out that there are no cookbook rules for policy judgment, and it is inevitable that policymakers are required to make a discretionary judgment. Policymakers are unlikely to design a proper policy response without full knowledge of the nature of asset price hikes or a correct forecast of potential growth rates. In any policy response, it is deemed important to assess financial and macroeconomic stability from the viewpoint of sustainability.
Reference


Fisher, I., 1933. The Debt Deflation Theroy of Great Depressions. *Econometrica*, 1,


Figure 1. The trends and growth rates of Housing Price Index (HPI), Consumer price index (CPI), and real GDP per capita of Taiwan (quarterly frequency)

(Source: Taiwan Economic Data Center (AREMOS))
Table 1. Summary Statistics of change rates of real GDP per capita, CPI and HPI

<table>
<thead>
<tr>
<th></th>
<th>ΔHPI</th>
<th>ΔCPI</th>
<th>ΔGDP</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Mean</strong></td>
<td>-0.0003</td>
<td>0.0043</td>
<td>0.0098</td>
</tr>
<tr>
<td><strong>Std. Dev.</strong></td>
<td>0.0246</td>
<td>0.0090</td>
<td>0.0258</td>
</tr>
<tr>
<td><strong>Skewness</strong></td>
<td>-0.1803</td>
<td>-0.2255</td>
<td>-0.0369</td>
</tr>
<tr>
<td><strong>Kurtosis</strong></td>
<td>3.2023</td>
<td>2.4020</td>
<td>2.1194</td>
</tr>
</tbody>
</table>

Notes: The base year is 2001. The level variables are in differential form after taking a log.

\[
\Delta \text{HPI} = \log(\text{HPI}_t / \text{HPI}_{t-1}), \quad \Delta \text{CPI} = \log(\text{CPI}_t / \text{CPI}_{t-1}), \quad \Delta \text{GDP} = \log(\text{GDP}_t / \text{GDP}_{t-1}).
\]
Table 2. Unit Root Tests

<table>
<thead>
<tr>
<th></th>
<th>HPI</th>
<th>CPI</th>
<th>GDP</th>
</tr>
</thead>
<tbody>
<tr>
<td>Level</td>
<td>-1.731(4)</td>
<td>-2.484(12)</td>
<td>-1.712(58)</td>
</tr>
<tr>
<td>First differential</td>
<td>-6.885***(2)</td>
<td>-8.805***(6)</td>
<td>-8.866***(13)</td>
</tr>
</tbody>
</table>

Notes: 1. Variables are formed in natural logarithm. The critical values for 1%, 5%, and 10% level of ADF and PP are -4.121; -3.487 and -3.172 respectively. The critical values for the ADF t-statistics are from the MacKinnon (1996) table.
2. The numbers in the parentheses [.] of ADF are the appropriate lag lengths selected by MAIC, whereas the numbers in the parentheses (.) of PP are the optimal bandwidths decided by the Bartlett kernel of Newey and West (1994).
3. *** and ** denote the significance at 1% and 5% level, respectively.
### Table 3. Zivot and Andrews Unit-Root Test with a Structural Break

<table>
<thead>
<tr>
<th>Model</th>
<th>Break</th>
<th>Z-A test</th>
<th>Break</th>
<th>Z-A test</th>
<th>Break</th>
<th>Z-A test</th>
</tr>
</thead>
</table>

Notes: 1. The critical value for 1%, 5%, and 10% levels are -5.57, -5.08, and –4.82 for Model C and -5.57, -5.08, and –4.82 for Model C/T from Zivot and Andrews (1992).
2. The numbers in the parentheses [.] are the appropriate lag lengths selected by MAIC.
Table 4. Determination of Cointegration Rank in Consideration of a Linear Trend and a Quadratic Trend

<table>
<thead>
<tr>
<th></th>
<th>Model 1</th>
<th>Model 2</th>
<th>Model 3</th>
<th>Model 4</th>
<th>Model 5</th>
</tr>
</thead>
<tbody>
<tr>
<td>Rank</td>
<td>H_2</td>
<td>H_1*</td>
<td>H_1</td>
<td>H*</td>
<td>H</td>
</tr>
<tr>
<td></td>
<td>T_2(r)</td>
<td>C_2(5%)</td>
<td>T_1*(r)</td>
<td>C_1*(5%)</td>
<td>T_1(r)</td>
</tr>
<tr>
<td>r = 0</td>
<td>49.15**</td>
<td>24.27</td>
<td>56.18**</td>
<td>35.19</td>
<td>44.54</td>
</tr>
</tbody>
</table>

Notes: 1. T_2(r), T_1*(r), T_1(r), T*(r), and T(r) denote the Trace-test statistics for all the nulls of H(r) versus the alternative of H(p) of Johansen’s five models.
2. C_2(5%), C_1*(5%), C_1(5%), C*(5%), and C(5%) are the 5% critical values for Johansen’s five models.
4. VAR length is 3 for all models, which is selected based on the minimum AIC.
<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Model</th>
<th>Break</th>
<th>G-H</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>C/T</td>
<td>1997:03</td>
<td>-3.914 [8]</td>
</tr>
</tbody>
</table>

Notes: 1. GH denotes the **Gregory and Hansen** test statistics. Critical values for model C, C/T and C/S are -4.92, -5.29, -5.50 respectively. (see: Gregory and Hansen (1996) Table 1. p.109)
2. The numbers in the parentheses [.] are the appropriate lag lengths selected by AIC.
### Table 6. Pairwise Granger Causality Test

<table>
<thead>
<tr>
<th>Independent variables</th>
<th>Lagged dependent variables</th>
</tr>
</thead>
<tbody>
<tr>
<td>{Break date}</td>
<td>ΔHPI_{t-i}</td>
</tr>
<tr>
<td>Model A</td>
<td></td>
</tr>
<tr>
<td>ΔHPI</td>
<td>-1.253</td>
</tr>
<tr>
<td>ΔCPI</td>
<td>0.066</td>
</tr>
<tr>
<td>ΔGDP</td>
<td>0.179</td>
</tr>
<tr>
<td>Model B (C)</td>
<td></td>
</tr>
<tr>
<td>ΔHPI</td>
<td>-1.706</td>
</tr>
<tr>
<td>ΔCPI</td>
<td>0.014</td>
</tr>
<tr>
<td>ΔGDP</td>
<td>0.222</td>
</tr>
<tr>
<td>Joint test</td>
<td></td>
</tr>
<tr>
<td></td>
<td>[1.578]*</td>
</tr>
<tr>
<td>Model C (C/T)</td>
<td></td>
</tr>
<tr>
<td>ΔHPI</td>
<td>-2.751</td>
</tr>
<tr>
<td></td>
<td>[8.235]**</td>
</tr>
<tr>
<td>ΔCPI</td>
<td>0.120</td>
</tr>
<tr>
<td></td>
<td>[6.824]*</td>
</tr>
<tr>
<td>ΔGDP</td>
<td>0.287</td>
</tr>
<tr>
<td>Joint test</td>
<td>503.91</td>
</tr>
<tr>
<td>Model D (C/S)</td>
<td></td>
</tr>
<tr>
<td>Pre-break Period</td>
<td></td>
</tr>
<tr>
<td>ΔHPI</td>
<td>-0.262</td>
</tr>
<tr>
<td>ΔCPI</td>
<td>0.059</td>
</tr>
<tr>
<td>ΔGDP</td>
<td>0.214</td>
</tr>
<tr>
<td>Post-break Period</td>
<td></td>
</tr>
<tr>
<td>ΔHPI</td>
<td>0.029</td>
</tr>
<tr>
<td>ΔCPI</td>
<td>0.022</td>
</tr>
<tr>
<td>ΔGDP</td>
<td>0.182</td>
</tr>
<tr>
<td>Joint test</td>
<td>497.78</td>
</tr>
</tbody>
</table>

Notes:
1. **, * denote the 5% and 10% significant level respectively.
2. The sums of coefficients and Chi-square Test statistics are reported in parentheses [ ]. The values in [ ] denote the Chi-square Test statistics of Joint test on coefficients of dummy variables. The values in [ ] present the Chi-square Test statistics of Joint test on time trend coefficients. The values in ( ) are the t test statistics on singular coefficient of dummy variable or of time trend.
3. Lag length is 3 selected by AIC.
Table 7. Causality Direction

<table>
<thead>
<tr>
<th>Model A: No dummy</th>
<th>Model B: Constant dummy (C)</th>
<th>Model C: Constant dummy + trend (C/T)</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \Delta CPI \leftrightarrow \Delta HPI )</td>
<td>( \Delta CPI \times \Delta HPI )</td>
<td>( \Delta CPI \quad \Delta CPI )</td>
</tr>
<tr>
<td>( \Delta GDP \times \Delta HPI )</td>
<td>( \Delta GDP \times \Delta HPI )</td>
<td>( \Delta GDP \times \Delta HPI )</td>
</tr>
</tbody>
</table>

Model D: Constant dummy + switch model (C/S)

<table>
<thead>
<tr>
<th>Pre-break Period</th>
<th>Post-break Period</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \Delta CPI \times \Delta HPI )</td>
<td>( \Delta CPI \times \Delta HPI )</td>
</tr>
<tr>
<td>( \Delta GDP \times \Delta HPI )</td>
<td>( \Delta GDP \times \Delta HPI )</td>
</tr>
<tr>
<td>( \Delta GDP \leftrightarrow \Delta CPI )</td>
<td>( \Delta GDP \leftrightarrow \Delta CPI )</td>
</tr>
</tbody>
</table>

Notes: The symbol \( \leftrightarrow \) denotes the bilateral causality with positive effect. \( \leftrightarrow \) and \( \rightarrow \) denote the unilateral causality with positive effect. \( \rightarrow \) denotes the unilateral causality with negative effect. And \( \times \) denotes the non-causality.