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Risk and Predictability of Singapore's Direct Residential Real Estate Market

by

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Abstract

This study explores the topic of the predictability of direct real estate prices in the short-run and the risks facing investors via a case study. Two models are estimated using heteroscedastic and autocorrelation robust ML method. Possible structural shifts of the models are examined. The one assuming that the model captures all the economic influences produces slightly better in-sample fitting. The other model assumes that there could be some important information which is not publicly available. Such information can nevertheless be extracted using Kalman filter. The latter has smaller forecast error in general. We found that a rational speculative bubble is an important predictor of short-run price movement, especially when the market is volatile and noisy. Rental is the only fundamental variable which has any important role to play in the short-run price generating process. Further more, the influence of rental is significant only when the market is inactive. Based on

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the study, we argue that the risk facing market participants comes not from the rational speculative bubble given its predictability, but primarily from unpredictable local policy shifts.

JEL classification: G12, C13, C52

Keywords: Risk; information; rational bubble; Kalman filter.

1. Introduction and Review of Literature

The objective of this study is to explore the role of economic fundamentals, the role of speculation and the risks involved in the private residential market of Singapore. The economists in Singapore have shown great interests in the real estate market of the country, as is evident by their research outputs in the last decades or so. Some of them explore the relationship between the real estate prices and the economic fundamentals such as GDP (Ho and Cuervo, 1999; Zhang, 2004); the role of bank lending in the market (Koh et al, 2004); some take a micro approach such as the spatio-temporal autoregressive modeling of the condominium price determination by Sun et al (2005); some study the legal issues (Phang and Wong, 1997; Phang 2005). Others look at the connection between the direct and the indirect real estate market. Studies on the risk characteristics of the market focus on the indirect real estate market (Clascock et al, 2006; Liow, 2005; Tien et al, 2004). The current study intends to fill the gap by directly measuring the relative importance of rational speculative bubble to the economic fundamentals in the short-run price determination, and by identifying the source of risks facing participants in Singapore's direct private residential market.

Singapore is a densely populated city state where on average each person has approximately 160 m² of land. According to Malpezzi and Wachter (2002), a city with restricted land supply is highly likely to see rampant speculative bubbles in the real estate market. However, the housing sector in Singapore is frequently intervened by the government, largely via its direct provision and tight regulation of the public housing, called HDB flats as they are built by the Housing and Development Board. Approximately 85% of housing stocks in Singapore are HDB flats. One of the objectives of these interventions is to prevent HDB dwellings from becoming a vehicle of speculation. Researchers have shown

that the policies designed to regulate the HDB housing sector have significant impact on the private housing market. The paths of the two housing price indices also suggest that the two sectors are integrated (Phang and Wong, 1997; Phang, 2005; Zhang, 2004). Given that the government has a strong interest in protecting its citizen's welfare and wealth, one would expect the desire to speculate in the relatively small private residential market be dampened. Hence changes in fundamental economic variables such as GDP and rental, rather than speculative motivation, should have more important impact on the private housing prices. At the same time, there are mixed evidence on the integration of the direct and indirect real estate market² (Tien et al, 2003). The direct real estate market is immobile, and the risks involved in it should be idiosyncratic; whereas the indirect real estate market is subject to the influence of the general stock market which is highly globalized nowadays (Liow, 2005). If the two are segmented, then random global shocks would not constitute much risk to the investor in the indirect market.

Our empirical study, however, shows that a rational speculative bubble is a key driver behind the short-run price movement in the market of concern. Rental is the only variable having non-negligible influence on the price movement in the short-run. This influence is significant only when the market is inactive. The path of the bubble is nevertheless predictable, hence presenting little risk to the investors. A major source of risk, in fact, comes from the unpredictability of the local policy shifts. Of those, changes in credit policies have particular profound impact. Global shocks, like Asian Financial Crisis, dot-com bubble burst also have left their finger prints. However, these external shocks are relatively much less frequent and their impact is limited.

² The indirect real estate market refers to the market trading real estate related securities, as opposed to real estates directly.

2. A Description of Singapore's Private Residential Market

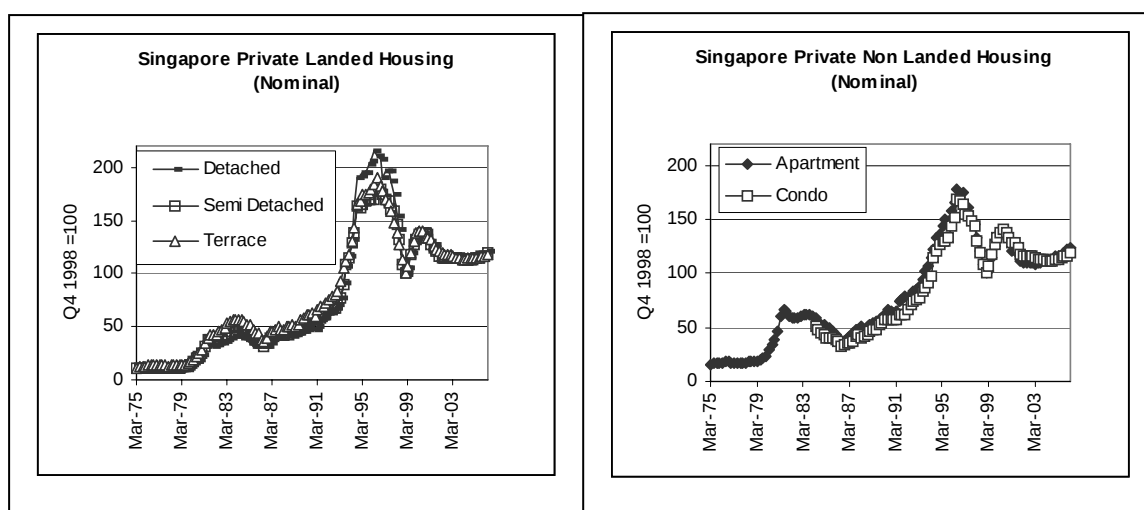
Singapore's private residential property market embarked on a ten-year-long rising trend in the late 1986. The duration, persistence and the magnitude of the rise dwarfed that experienced by its counterpart in Hong Kong (Xiao, 2006). During that time, every category in this market — detached, semi detached, terrace, apartment and condominium — gained significant value. The gain was measured before 1993, but started to go wild thereafter. Taking for example the condominium, the “modest” of all, more than quadrupled in real terms during the ten-year-long rally, more than two-thirds of the gain came between 1993Q1 and 1996Q2. The “star” performer, the detached houses, rose nearly by 600% in real value, almost three-fourth of which was gathered in the last three years and a quarter (table 1 and figure 1). In fact, the steepest rise of this category occurred between 1993Q1 and 1994Q4. The slope of its price path almost approached infinite between 1994Q3 and 1994Q4, resembling the “critical phenomenon” described by Sornette (2003)³. During this short time span, the averaged per annum increase of the price was 86.69% in real terms. The steep rise was followed by a steep fall in mid-1996. These observations arouse the curiosity of the authors: how much of the price growth (positive or negative) is due to fundamental change; how much due to market overreaction to fundamental changes; how much due to rational speculation; and how much neither the fundamental nor the rational speculation can explain? The unexplained part, if random, would measure the unpredictability hence riskiness of the market. We will also try to explore qualitatively the sources of such uncertainties.

Table 1 the Change in CPI Deflated Private Residential Price between June 1986 and June 1996

³ Critical phenomenon: at a critical point, the sensitivity of the price with respect to a small system perturbation goes to infinity. A critical point signals a looming market crash, due to system instability (Sornette, 2003).

	Detached	Semi Detached	Terrace	Apartment	Condominium
Jun-86	31	32.2	36.8	33.9	32.9
Mar-93	76.4	77.7	92.7	87.9	77.1
Gain since Jun-86	146.45%	141.30%	151.90%	159.29%	134.35%
Jun-96	211.3	178	189.8	178	169.1
Total gain since Jun-86	581.61%	452.80%	415.76%	425.07%	413.98%

Figure 1 the Path of Private Residential Prices by Category. Nominal Values are Used Due to the Lack of Data on CPI before January 1983.



Source: CEIC database

3. Theoretic Framework

This section presents a variant of the present value model. The model will serve as a theoretical foundation for the empirical stochastic model to be built in the next section. If

economic agents are risk neutral, the price of property will be equal to the expected discounted present value of the rent accruing to ownership of the property during the ownership period, plus the price at which the property can be sold at the end of the ownership period. There could be other economic variable that has some influence on the price via their impact on expectation formation. Mathematically,

$$P_t = \frac{E_t[P_{t+1} + D_t + C_t]}{1 + R_t}, \quad (1)$$

where P_t = the real price of the property asset at time t ; D_t = the real rent received during the period t ; R_t = the time varying real discount rate; and C_t = other relevant economic variables.

Define

$$r_t \equiv \log(1 + R_t); \quad (2)$$

hence,

$$r_t \equiv \log(E_t[P_{t+1} + D_t + C_t]) - \log(P_t). \quad (3)$$

In a static world, rent grows at a constant rate, g , and the log of rent to price ratio is also a constant. In such a world, the log of the gross discount rate, r_t , would also be a constant. If the transversality condition, $\lim_{i \rightarrow \infty} \rho^i E_t[p_{t+i}] = 0$, is satisfied, we would have the fundamental solution for the price:

$$p_t = p_t^f = \frac{\kappa - \xi}{1 - \rho} + (1 - \rho) \sum_{j=0}^{\infty} \rho^j E_t[d_{t+j} + c_{t+j}]. \quad (4)$$

where p = the log of P ; d = the log of D ; and c = the log of C .

However, the transversality condition may fail to hold. In such a case, we would expect the price to contain a *rational speculative bubble*, b

$$p_t = p_t^f + b_t, \quad (5)$$

Where the bubble component

$$E_t[b_{t+i}] = \frac{1}{\rho^i} b_t. \quad (6)$$

This implies that

$$E_t[b_{t+1}] = \frac{1}{\rho} b_t. \quad (7)$$

If the log of the property prices and the log of the rents are I(1) process, the following model may be estimated instead:

$$\begin{aligned} \Delta p_t^f &= p_t^f - p_{t-1}^f \\ &= (1-\rho) \sum_{j=0}^{\infty} \rho^j \{ [E_t[d_{t+j}] - E_{t-1}[d_{t+j-1}]] + [E_t[c_{t+j}] - E_{t-1}[c_{t+j-1}]] \}. \end{aligned} \quad (8)$$

Suppose the growth of the rents and the other relevant economic variables follow respectively $AR(p)$ and $AR(q)$,

Then

$$\Delta p_t^f = \sum_{i=0}^p \theta_i \Delta d_{t-i} + \sum_{j=0}^q \psi_j \Delta c_{t-j}. \quad (9)$$

Where θ and ψ are functions of the coefficients of the underlying processes of the price, the rent, the other relevant variables.

If a bubble is present,

$$\Delta p_t = \Delta p_t^f + \Delta b_t, \quad (10)$$

and

$$E_t[\Delta b_{t+1}] = \frac{1}{\rho} \Delta b_t. \quad (11)$$

Suppose there is some private information unavailable to the researcher, hence not included in equation (10). Let s_t denotes a model specification error, so

$$\Delta p_t = s_t + \sum_{i=0}^p \theta_i \Delta d_{t-i} + \sum_{j=0}^q \psi_j \Delta c_{t-j} + \Delta b_t \quad (12)$$

And

$$\Delta s_{t+1} = \beta \Delta s_t \quad (13)$$

This error is unobserved but can be inferred using Kalman filter (refer to Harvey, 1989; Hamilton, 1994 for the technical details of this filter).

4. Econometric Examinations

a. Preliminary data analysis

A property is a durable consumption good as well as an investment asset. In the property price literature, there is no unified theoretical framework which clearly defines the property price fundamentals. Typically, different researchers will use different fundamentals in their theoretical or empirical models (Ito and Iwaisako, 1995; Clayton, 1996; Ho, 1999 Kim and Lee, 2000; Hendershott, 2000; Basile and Joyce, 2001). The most prominent asset price model in the financial economics literature is the present value model. Based on that model, the faster the rent grows the faster the price rises. This is because that rent is the return to the capital tied up in the property. There could be other economic variables which are fundamental to property prices, such as GDP, wage, unemployment rate, mortgage rate, etc. In general, higher GDP and wage growth rate, lower unemployment rate boost buyers' confidence about their future earning power, hence they are more willing to take out

mortgages. They also increase the expected capital gains from entering the property market, hence encourage buying. Mortgage rate, on the other hand, measures the cost of capital. The higher the cost of the capital, the lower the incentive one has to buy a property. Vacancy/occupancy rate may also play a role. The loan to residential property has been shown to have substantial explanatory power, but it is regarded as a bubble generator rather than a fundamental force (Ito, 1995; Krugman, 1998; Koh et al., 2004). This variable will not be included in the model to be built, as we have an alternative bubble proxy which is most readily available, and the inclusion of both could result in multicollinearity. The relationship between the price and these variables may be nonlinear, but we will assume that it is log linear in this paper, based on the theoretical model in section 3 of the paper.

Since the rental indices are not available for individual categories before December 1998, we have to compromise to use the private residential property price index, which is the weighted average of price indices of the various housing categories. This price index is available only quarterly. Therefore, all the other variables are also transformed accordingly. The prime lending rate is used as an indication of the mortgage rate offered in the market; and average monthly earnings as an indication of wage. All data are taken from CEIC database⁴.

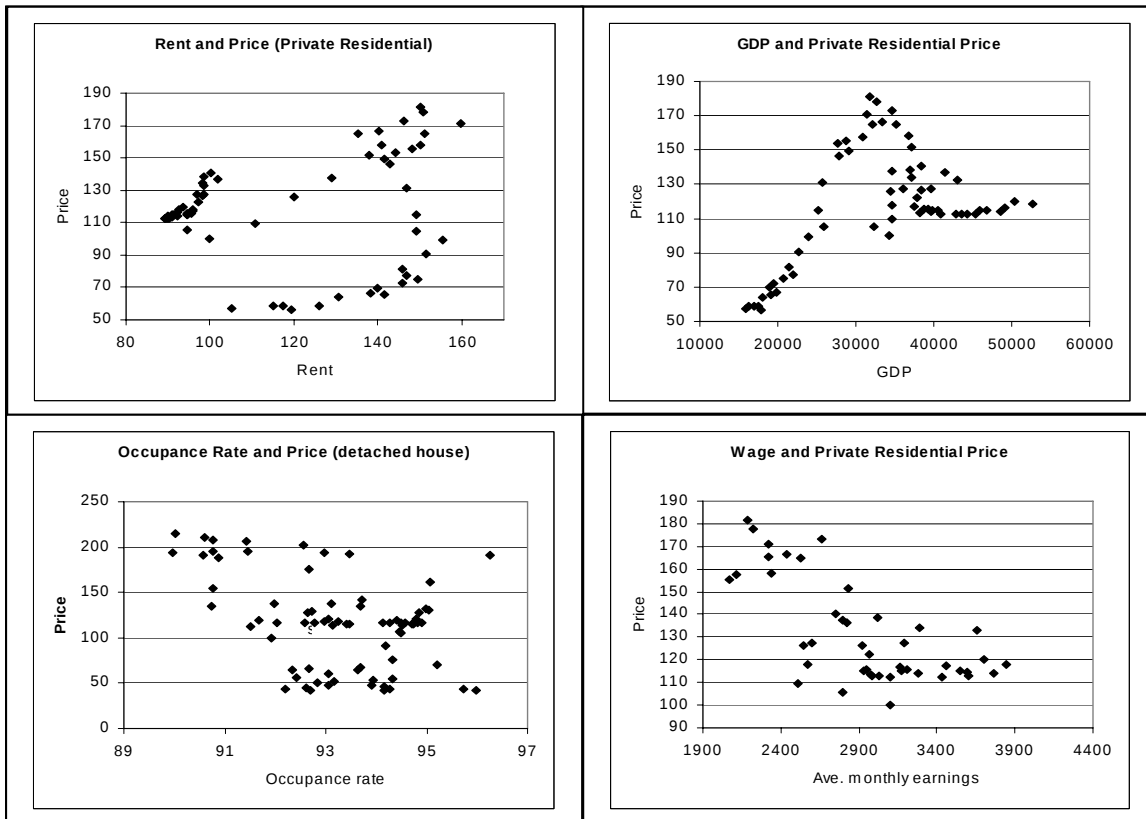
The preliminary x-plot shows that there is no identifiable short-run relationship of the price index with the occupancy rate and with the prime lending rate⁵; a weird negative correlation of the price with the average monthly earning; an expected positive correlation with the GDP, and with the residential property rental indices; and a weak negative correlation with the

⁴ Refer to <http://www.ceicdata.com/>

⁵ Ho and Cuervo (1999) show that not even long-run relationship exist between this price and this prime lending rent. Also refer to Phang and Wong (1997) for their conclusion on the role of various economic variables and policies in this market.

unemployment rate. But the correlations between the price and the rent and that between the price and GDP shifted at some point, suggesting structural changes (figure 2).

Figure 2 X-plot of the Price and Some Variables (Nominal)



The variables showing some relationship with the price, weird or as expected, are selected together with the price for further investigation. These variables are unemployment rate, GDP, rent, average monthly earning. Price, GDP, rent and average monthly earning are deflated using CPI.

The null hypothesis that the Pearson correlation coefficients equals to zero can be rejected for all except for that between the price and the rental indices when the level variables are considered. When we take log transformation of the real price, GDP, rent and wage and re-

conduct the test, the null hypothesis can be rejected for all except for the correlation coefficient of the price with the rent and with the unemployment rate. Note that throughout this paper we will use 5% significance level for various tests taken, unless mentioned otherwise.

To avoid type II spurious regression, the Philips-Perron unit root test is applied to the level and log transformed variables. The test does not reject the null hypothesis of unit root when the above variables, in level and in log, are considered. We therefore take the first difference of the logs (except for unemployment, where the difference is taken on the level), and this time the null hypothesis can be rejected. The interpretation of the differenced logs is straightforward: the growth rate of the corresponding variables.

We choose to difference the logs rather than the level values because of the log-linearity assumption mentioned above. In addition, we intend to explicitly estimate a rational speculative bubble in this study. A log transformation will allow for negative bubble without the funny result of negative price⁶. Furthermore, we will make use of the Kalman filter in this study to extract unobservable market information. The use of this algorithm requires that certain variables are normally distributed. When the Shapiro-Wilk test is applied to the level and the log values, the null hypothesis of normality can be rejected for all variables considered, except for the average monthly earning. It *cannot* be rejected however, when the test is applied to the first difference of the log price and rent. Many researchers agree that Shapiro-Wilk W test is the most reliable normality test for small and medium sized samples (Conover, 1999; Shapiro and Wilk, 1965; Royston, 1982a, b, 1995). This result is consistent with the general wisdom that the distribution of individual stock returns are not normal, however, the returns to stock indices averaged over long period such as a month are close to

⁶ Refer to Xiao (2005) for argument of existence of negative bubble.

normal (Campbell et al, 1997; Drozd, S. also made similar comments in Econophysics Colloquium 2006 held in Tokyo).

We apply the correlation test to the concurrent transformed variables. The probability is 0.1519, 0.1068, 0.0012 and 0.5507 respectively that there is zero correlation of the price with the unemployment rate, GDP, rental and average monthly earning. When the lagged differenced variables are included, the null hypothesis cannot be rejected for all pairs excepted for the price and its own lagged value.

Table 2 Pearson Correlation Coefficients, $|r|$, between Price (p_t) and other variables. Variables are in the first difference of log real values, except for unemployment rate, which is in the first difference of the level value.

	Unemployment rate (u)	GDP (gdp)	Rental (r)	Average monthly earning (w)	Price lagged one period (p_{t-1})
$ r $	-0.1812	0.2035	0.3965	-0.0936	0.7287
$Pr > r $	0.1519	0.1068	0.0012	0.5507	<.0001

These preliminary tests suggest that, to explain the short-run variation of the price, we may want to include the GDP, rent and lagged price (all in differenced logs) as regressors. However, the regression coefficient on GDP has the wrong sign, which could have been caused by multicollinearity. We replace it by unemployment rate (in differenced value), since the two are significantly negatively correlated ($Pr > |r| = 0.3579$). The regression coefficient of this variable has the expected negative sign.

The unemployment rate and rental are intuitively fundamental to the housing prices, but what about the lagged prices? According to Shiller (1990), it is an indication of speculative demand. If price growth drives further price growth, it implies a rational speculative bubble is building up. Hence, the lagged price functions as a proxy for such bubble in the model to be built. Another possible bubble proxy is housing related loans. For two reasons, however, this is not selected. The causation between housing price and housing loan could be two-way and the current approach may not be suitable for modeling that relationship (Bernanke and Gertler, 1989; Kiyotaki and Moore, 1997; Bernanke et al, 1998; Gerlach and Peng, 2004; Goodhart, 1995). Furthermore, the lagged price is more readily available as a bubble proxy for most researchers.

b. Model Building, Structural-break Identification and Methodology Justification

Based on the theoretical framework and data analysis, we attempt to build one empirical stochastic model including the concurrent and the lagged unemployment rate, rental and the lagged price as regressors, and another model which will in addition includes a latent state variable regressor. This variable is meant to capture all the other information which is not available to the general public. We will refer to the former as the no-S model and the later the SS model. The number of lags is selected using F test by stepwise regression. The criteria used for model selection are the usual AIC and SBC. The selected models are show below

No-S Model:

$$\Delta p_t = \theta_1 \Delta U_t + \theta_2 \Delta r_t + \theta_3 \Delta p_{t-1} + \varepsilon_t \quad (14)$$

with $E[\varepsilon_t] = 0 \quad Var[\varepsilon_t] = \sigma^2$

where Δp_t is first difference of the log price; ΔU_t the change in unemployment rate; Δr_t the first difference in log rental; and ε_t a random price innovation which measures the risks facing market participants. The subscript t indicates the time of the observation.

SS Model:

$$\begin{aligned} y_t &= s_t + x_t' \theta + \varepsilon_t, \text{ with } \varepsilon_t \xrightarrow{d} N(0, \Sigma) \\ s_t &= \beta s_{t-1} + v_t \text{ with } v_t \xrightarrow{d} N(0, V) \\ \text{and } E(\varepsilon_t, v_s) &= 0 \text{ for all } s, t \end{aligned} \quad (15)$$

where y_t is the dependent variable, the growth rate of the housing price, x_t the same observable explanatory variables as those in the no-S model, and s_t a latent state variable capturing all the other unobserved fundamentals influencing the price. The parameters in the model are time-invariant. The state variable can be inferred using Kalman filter (Harvey, 1989). Both models are estimated using heteroscedasticity and autocorrelation robust maximum likelihood method (SAS/ETS, version 8, ch.8; SAS/IML, version 8, ch.10).

As the x-plots suggest structural changes, we estimate the model using the whole sample as well as the sub-samples before and after a pre-specified breaking point. Due to the lack of data on the variables used as regressors, the whole sample span from Mar 1990 to March 2006 (quarterly). Four candidate breaking points are examined. They are 1996Q2, 1999Q2, 1999Q3 and 2000Q3. The rational of the choice will be explained below.

At first, the breaking point was chosen at 1996Q2, as we know that there was a policy intervention at that time to curb the market speculation (Phang, 2005 p.14). Another reason for choosing 1996Q2 is that the price plunged into a free fall from this point onwards after a ten-year long rally. Analogous to Zhandi's (1999) argument, the response of the price to

changes in the fundamentals and in the rational bubble when the price is falling may differ from that when the price is rising.

The Chow test shows that the probability of no structural change at this point is 0.4020. The sub-sample regression result is rather poor as well, both in terms of in-sample fitting and out-of-sample forecast. The R-square for the whole sample is 0.5425/0.5378, with a one-step squared forecast error of $4.6 \times 10^{-5}/1.1 \times 10^{-4}$ when no-S/SS model is fitted. The corresponding figures for the before-break sub-sample are 0.2885/0.1849 and $7.1 \times 10^{-4}/2.3 \times 10^{-3}$ respectively; and those for the after-break sample are 0.4969/0.4955 and $5.9 \times 10^{-5}/2.3 \times 10^{-5}$ (ΔU_t is replaced by ΔU_{t-1} as regressor in this sub-sample as will be explained below).

We therefore attempt an alternative procedure based on residual volatility to identify the breaking point(s). First, we estimate the SS model using the whole sample. The standardized residuals (defined in Harvey, 1989 p257), which should be approximately distributed $NID(0, 1)$ when the model is correctly specified, are exported for diagnostics. The plot of these residuals shows a clear reduction in volatility since 1999Q3. Only one such residual (2000Q3) falls outside one standard deviation of the theoretical distribution after 1999Q3, whereas plenty of these do so before this point. Ex post, this is consistent with the conclusion reached by Zhang (2004) using recursive segmentation method. Zhang identified 1999Q2 as the breaking point for this same price index (Q1 1990 to Q1 2004)⁷.

Hence, we also consider 1999Q2 and 2000Q3 as possible choice of breaking point(s). The theoretical probabilities of no-break are respectively 0.1049, 0.1278 and 0.9410 for 1999Q2, 1999Q3 and 2000Q3. The stepwise regression is carried out again for all the sub-samples.

⁷ Zhang (2004) has a different model specification, with more regressors including some of those we have considered in this paper.

The selected regressors for each before-break sample are the same as those in equation (1) and (2). For the after-break sample, the concurrent unemployment is replaced by the variable lagged one period. The sub-samples are re-estimated using maximum likelihood method. We settled down at one breaking point due to the limited observations. That point is 1999Q3. The choice is based on the probability of structural change as well as on balancing between out-of sample forecast performance and in-sample goodness-of-fit (table 3).

Figure 3 Standardized Residuals from SS Model with the Whole Sample

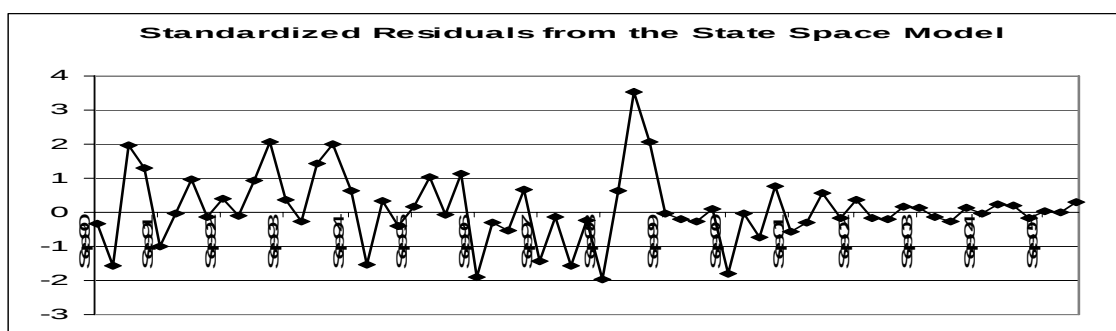


Table 3 Comparison of Different Breaking Points: R-square and one-step squared forecast error

		R ²		Squared forecast error	
		SS	No-S	SS	No-S
1996Q2	Before	0.1849	0.2885	2.3 10 ⁻³	7.1 10 ⁻⁴
	After	0.4955	0.4969	2.3 10 ⁻⁵	5.9 10 ⁻⁵
1999Q2	Before	0.5474	0.5690	4.1 10 ⁻³	3.2 10 ⁻³
	After	0.7552	0.7573	5.5 10 ⁻⁷	8.2 10 ⁻⁶
1999Q3	Before	0.5594	0.5800	3.0 10 ⁻⁵	9.3 10 ⁻⁵

	After	0.6388	0.6448	$1.1 \cdot 10^{-7}$	$9.9 \cdot 10^{-6}$
2000Q3	Before	0.5127	0.5199	$1.2 \cdot 10^{-2}$	$1.6 \cdot 10^{-2}$
	After	0.6210	0.6213	$9.5 \cdot 10^{-6}$	$8.8 \cdot 10^{-6}$
No breaking		0.5378	0.5425	$1.1 \cdot 10^{-4}$	$4.6 \cdot 10^{-5}$

The Chow test shows that the theoretical probability of no change at 1999Q3 is 0.1278. Bear in mind that, with unequal variances, the actual probability of observing the value of this test statistic is less than the theoretical one. The Wald test was insignificant when the model is estimated without the latent state variable (Wald = 5.83 with 3 degree of freedom). It is significant at 1% when this variable is incorporated and estimated jointly with the model (Wald = 37.39 with 4 degree of freedom). Kim and Perron (2006) show that, for structural break, Wald-based test is superior to LR and LM tests. Greene (1997) also mentions that, with large sample and unequal variances, a Wald test is preferred. However, with small or moderately sized samples the actual probability of the test statistic is larger than the critical value one use to carry it out (Greene, 1997). In our case, the effective sample size is 63 which can be considered large by the usual criterion. To be sure about the exact probability, one might want to calculate it using simulation. However, for our purpose, the above tests will suffice.

The use of Kalman filter relies on the normality assumptions on disturbances and the initial state vector. It is weakly justified in the current context on the following ground: at 5% significance level, the null hypothesis of normality cannot be rejected for the price and rent in the whole sample and the sub-samples, for the unemployment rate in the sub-samples, and for the residuals from the before-break sample. At 1% level, this hypothesis cannot be rejected for the residuals from the after-break sample (table 4). The initial state is not

observable, and is assumed normal using the same argument given in section 4a. That is, daily observations of a variable may be far from being normal, but quarterly observations are close to normal. The normality tests have also been applied to the other considered sub-samples, and the results are largely consistent.

Table 4 Shapiro-Wilk Test of the Null Hypothesis: Variables are Normal

		Whole sample	Before 1999Q3	After 1999Q3
W Pr < W	Price	0.9803	0.9690	0.9563
		0.3967	0.3659	0.3237
	Rent	0.9725	0.9879	0.9622
		0.1623	0.9488	0.4360
	Unemployment	0.9282*	0.9519	0.9419
		0.0011	0.1027	0.1494
	Residuals	0.9412*	0.9913	0.9046**
		0.0047	0.9911	0.0198

Note: * means significant at 1%, ** significant at 5%. *** significant at 10%. The rest of tables in this paper will use the same notations. The test result for the whole sample is included for reference and comparison only.

c. Model Evaluation, Coefficients Interpretation and Diagnostics

The model without assuming the state variable has slightly better in-sample fitting for all the samples considered. But the one with the state variable as regressor generates far smaller forecast error in the sub-samples (table 3). This conclusion applies in five of the eight sub-

samples when we consider four different possible breaking points (1996Q2, 1999Q2, 1999Q3 and 2000Q3).

The R-square given in table 3 measures the average fit of the model to the data, which are in the range of 0.55 to 0.64 for the selected sub-samples. The following plot shows how well individual data point fare. The model predicts the market movement in the wrong direction 3 times before 1999Q3 and one time after 1999Q3. With reference to Phang and Wong (1997) and Phang (2005), the sources of some of these surprises to the model are given in table 5.

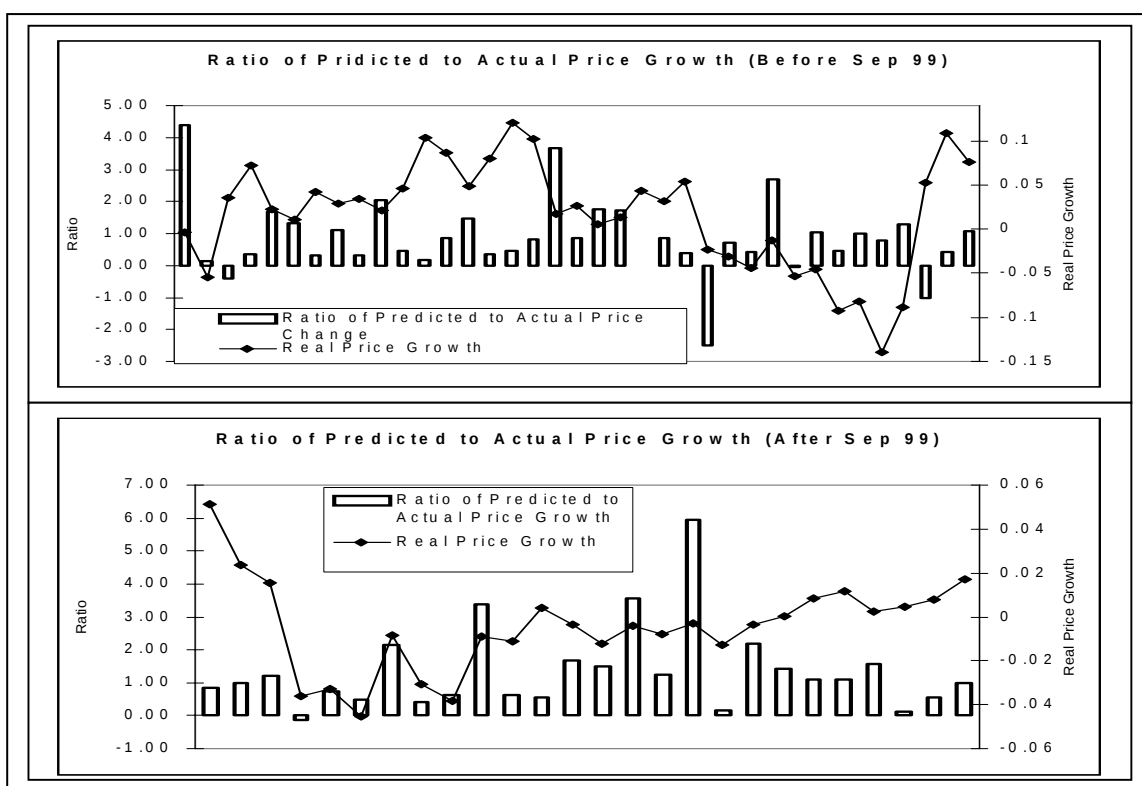
To take for example, in March 1993, the CPF⁸ board announced that it would, with effect from October 1993, allow withdrawals of CPF savings to be used to meet interest payments on mortgage loans for resale HDB and private housing purchases. Before this, members of the fund were allowed to withdraw only up to 100 per cent of the value of the property at the time of purchase. In April 1993, subsidized mortgage loans provided by the Housing and Development Board for purchase of resale HDB flats were increased to 80 per cent of the market value or resale price, whichever is lower. Previously, the HDB mortgage loan was based on up to 80 per cent of HDB's 1984 sale price for a comparable new flat. The market responded to these credit policies with a large jump in price growth, rendering a severe under prediction of the model in 1993Q2 (the ratio of predicted to actual price growth is 0.17). On 15 May 1996, the government introduced a anti-speculation package to cool the real estate market. These includes taxed on capital gains from the sale of any property within three years of purchase, stamp duty on every sale of property, limitation of housing loans, etc.

⁸ CPF stands for Central Provident Fund. The fund is essentially a pay-as-you-go social security scheme which requires mandatory contributions from both the employer and the employee. The scheme covers about two-thirds of labor force. Between 1968 and 1981, the fund could be withdrawn only for the down-payment, stamp duties, mortgage and interest payments incurred for the purchase of HDB flats. In 1981, the scheme was extended to allow for withdrawals for purchase of private housing (Phang and Wong, 1997).

The market took a nose dive in the following months. The model is caught by surprise and produced a large error (the ratio is -2.51).

Phang and Wong (1997) show that, in response to a policy shock, the initial increase/decrease in the private residential housing price is usually large. But the impact of the shock tends to dampen off over time. Most of the over- or under- predictions of the model look like temporary excess adjustment to the random shocks. The model will pick up the shock and adjust only with a lag, hence, the under- and over-prediction at the time of shock and afterwards.

Figure 4 Ratio of Predicted to Actual Price Growth



Note: The prediction would be perfect if the ratio of the predicted to the actual price growth is +1.

Table 5 Random Shocks to the Model

Time of severe under- or wrong prediction	Ratio of predicted to actual price growth	Events (Phang and Wong, 1997; Phang, 2005)
1993Q2	0.17	CPF liberalization (Mar. 93); HDB resale flat credit policy (Apr. 93).
1994Q1 and 1995Q4	0.36 and 0.01	Housing grants to first time purchaser of HDB resale flat.
1996Q3	-2.51	Anti-speculative package (May 96), including capital gain taxes, stamp duty, limitation of housing loans, and limiting foreigners to non S\$ denominated housing loans.
1997Q3	-0.03	Asian Financial Crisis
2000Q3	-0.13	Burst of dot-com bubbles

When the whole sample is used for estimation, only the coefficient on the lagged price is significant at the conventional levels (1%, 5%, and 10%). When we fit the model using the selected sub-sample before 1999Q3, the highly volatile period, only the coefficients on the unemployment rate and the lagged price are significant. In the quieter period after 1999Q3, rent starts to exert its influence alongside the other two variables. Although statistically significant, the impact of the unemployment on the price is minor. When the unemployment increases by 1%, the price drops on average by less than 0.05% in all sample periods considered. During the volatile time, if the rent increases by 1%, the price increases by less than 0.12%, which is insignificantly different from zero given the size of the coefficient

standard error. But if the price in the previous period rose by 1%, the current price will rise on average by about 0.65%⁹. In the quieter period, the price becomes more responsive to the rent variations. A 1% change in rent will result in an about 0.60% change in the price in the same direction, while the price response to its previous change drop to less than 0.45% (table 6). These observations indicate that, when the market is volatile hence information noisy, speculative motivation is the dominant driving force behind the short-run price variations. However, when the market is quiet, investors start to cool down and become more concerned with the rental returns of their investment. Nevertheless, even during this time, rational speculation is still an important driving force in the market. These conclusions are robust with respect to choice of different breaking points, although the exact magnitude of the response will vary.

Table 6 Parameter Estimates and Test of Joint Significance

		Unemployment (concurrent or lagged as is appropriate)	rent	Lagged price	β	Wald (d.f.)	Predictive F test F (d.f.)
Parameter	SS	-0.00725	0.1031	0.6992*	0.0010	18.04*	0.10
s.e.		0.0069	0.0875	0.0522	0.1571	(4)	(1,62)
(whole	no-	-0.0074	0.1405	0.6777*	-	18.75*	0.04
sample)	S	0.0101	0.1581	0.0998		(3)	(1,58)
Parameter	SS	-0.0455*	0.0271	0.6991*	0.0012	12.35**	0.02
s.e.		0.0165	0.0907	0.0568	0.1610	(4)	(1,36)

⁹ This coefficient is 0.83 if the model is fitted to the rising phase of the price only. This suggests that there could be multiple breaking points. We allow for one breaking point in the current study due to the sample size restriction.

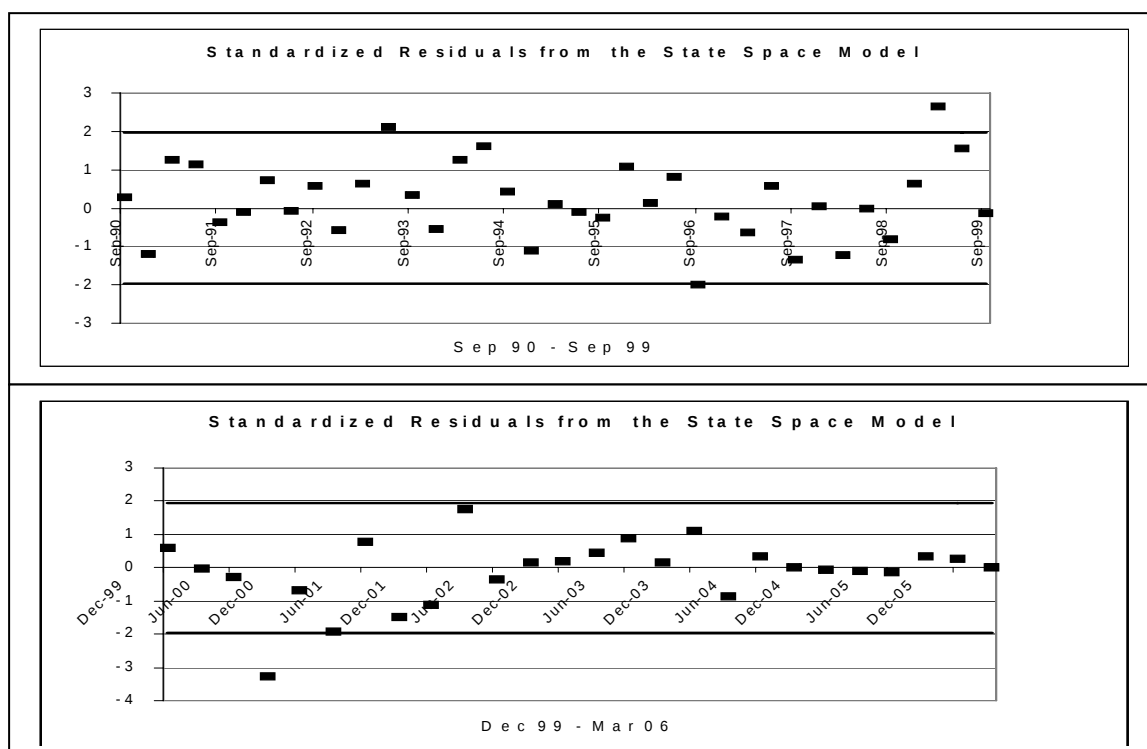
(before 1999Q3)	no- S	-0.0488** 0.0239	0.1166 0.2019	0.6339* 0.1326	-	8.36** (3)	0.05 (1,32)
Parameter s.e.	SS	-0.0090** 0.0033	0.6260* 0.1000	0.4416* 0.0433	0.0003 0.1626	29.27* (4)	0.00 (1,25)
(after 1999Q3)	no- S	-0.0088*** 0.0048	0.5872** 0.2587	0.4467* 0.1217	-	12.86* (3)	0.05 (1,21)

Note: (i) The coefficient of the state variable in the price equation is forced to be unity. This is because the sensitivity of price with respect to this variable is not important from policy point of view. It is the total impact of this variable on the price that matters; (ii) the null hypothesis of the Wald test is that the parameters are jointly zero. This is rejected for all; (iii) the null hypothesis of the F predictive test asserts that the model is correctly specified. This is not rejected in all cases. (iv) the whole sample is included for comparison and reference only.

In diagnostics using residuals from the no-S model, we do not reject the null hypothesis of homoscedasticity in the sub-samples Using LM test. When Durbin h test is applied to identify residual autocorrelation, the probability of no autocorrelation is 0.1478 for the before-break sample, and 0.4955 for the after-break sample. With residuals from the SS model, the homoscedasticity hypothesis cannot be rejected in the before-break sample. The LM test statistic is insignificant at 5% ($\chi^2(4) = 6.87$). This statistic is significant at 5% and 1% ($\chi^2(4) = 15.36$), for the other sub-sample. This is confirmed by the standard plots. The standardized residuals from the sample before the break are evenly spread about zero. All except three regression errors are within the 95% confidence interval. That from the after-break sample shows narrower spread towards the end of the sample, although only one is outside the 95% confidence interval. We tried to allow for GARCH(1,1) disturbances, but the AIC and SBC values suggest that the GARCH model is inferior. The forecast error of the model is also

larger. One way of mitigating this problem is to allow for one more structural break at around Dec 02. This is, however, not a very good choice given the limited observations we have. The other possible choice is to allow the coefficients to be a function of, say, the innovations in the previous period; or to shift value according to a Markov-switching process. These are, however, beyond the scope of this study.

Figure 5 Diagnostics with Standard Plot



The residual diagnostics suggest that the model residuals are random, indicating their unpredictability. The standard deviation of these residuals hence measures the riskiness of the market. Unsurprisingly, the period before 1999Q3 is more than three times riskier (estimated standard deviation = 0.0417) than the period afterwards (est. s.e. = 0.0134). The randomness of the residuals is a result of the random policy shifts and external shocks as argued above. Compared to the local policy shifts, the external shocks are much less

frequent and their impact is much smaller (table 5). Hence, the study indicates that a major source of risk facing the investors in this market is the unpredictable domestic policy shifts, especially those concerning the credit/loan availabilities. Although it does not have a mechanism to predict these shocks, each of the two models we built is able to incorporate immediately the new information carried by the innovation to correct or to reduce the prediction error (figure 4).

5. Conclusion and Discussion

In this empirical study, we explore the forces driving the private residential prices of Singapore, the predictability of the prices and how risky the market is between 1990 and early 2006. We model the price generating mechanism in two ways. In the first, we assume that the price growth is driven by publicly available information on economic fundamentals plus a rational speculative bubble (referred to as no-S model). In the second, we assume that some economic variables may not be observable by the researchers but nevertheless important (the SS model). These variables are lumped under a state variable, which can be extracted using Kalman filter. The use of this filter in this context is justified in the empirical section. Both models are estimated using heteroscedasticity and autocorrelation robust maximum likelihood method.

Given the established wisdom that there could be asymmetric responses of the price to the incremental changes in the explanatory variables, we allow for a structural break based on volatility. The selected breaking point, using plot of standardized residuals from the SS model, is of one data point difference from that selected by Zhang (2004). One of the two resulting sub-samples runs from 1990Q1 to 1999Q2; and the other from 1999Q3 to 2006Q1. We have

considered GARCH disturbances as an alternative approach to structural breaks. But GARCH models are inferior to our models by the AIC and SBC standard.

The no-S model produces slighter better in-sample fit, but the SS model generates much smaller out-of-sample forecast errors in most cases considered. This improved extrapolative forecast performance could be attributable to the possible fact that the state variable carries information not reflected by the publicly available economic variables. Either model shows that the public information, such as GDP, unemployment rate, prime lending rate, property occupancy rate, hardly has any impact on the short-run price movement. A rational speculative bubble dominated the scene when the market is volatile and noisy (before 1999Q3). Rental, a measure of the return to the capital tied up in a property, is a key determinant when the market is very quiet (after 1999Q3), but had hardly any significant influence on the price movement in the volatile period. Although taken a second place in the tranquil period, rational speculation is still far important than the fundamentals other than the rental.

Our models can explain 55% - 64% of the price growth. The models residuals are random, indicating their unpredictability. Examination of the individual predictions and the economic background show that the unexplained portion of the price change is due shocks external to the price generating mechanism. These shocks primarily come from the local policy shifts. This is because that the direct real estate market is immobile, and there is little evidence on the integration of this market with the mobile indirect real estate market (Tien and Shook, 2003). As a result, global shocks are far less important than the indigenous policy shocks. If so, the unpredictability of the policy path is an important source of risk in this market.

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