U.S. House Prices over the Last 30 Years: Bubbles, Regime Shifts and Market (In)Efficiency

Rose Neng Lai*

University of Macau

Robert Van Order

George Washington University

DRAFT

November 2014

Abstract

This paper studies the evolution of U.S. house prices across 45 metropolitan areas from 1980-2012. It uses a version of the Gordon dividend discount model, modelling price as present value of imputed rents as a measure of "fundamentals." This allows for a parsimonious specification, using only lagged rents, property values and interest rates (real and nominal) to explain property values. We find that cities share long run fundamentals, but adjust to them slowly — at a rate of around 10% per year, which is relatively constant across cities. However we also find sharp differences in short run adjustments across cities, which are correlated with local supply elasticities. Analysis of residuals suggests strong cyclical deviations from fundamentals throughout the period, with a high degree of serial correlation. The bubble period (2000-2007) was longer than usual and appears to have been extended after 2002, when it was dying out, in a way that is coincident with the rise in subprime securitization.

Keywords: Housing Bubbles; Pooled Mean Group Estimation; Supply Elasticities

JEL Classification: C23, R31

* Rose Neng Lai is Professor, Department of Finance and Business Economics, University of Macau, Taipa, Macau, China. e-mail: RoseLai@umac.mo. Robert Van Order is Oliver Carr Professor of Real Estate and Professor of Finance and Economics at George Washington University, Washington D.C.. e-mail: rvo@gwu.edu.

Introduction

The U.S. real estate market underwent a boom period from the end of the 1990s until about 2006 when property prices started to fall and default started to rise, leading to the collapse of the securitized mortgage, and then finally the collapse of many financial institutions. Studies about the housing bubble by now have become abundant. Examples are Black et al. (2006), Chan et al. (2001), Chang et al. (2005), Coleman et al (2008), Hwang et al. (2006), Taipalus (2006), and Wheaton and Nechayef (2008).

It has been 8 years since property values began to decline. A question is whether the property market is now under a new regime after the bubble burst, or instead is wandering back to the state before the bubble. If so, a follow up question would be whether there is yet a third regime or a long run phenomenon with which the bubble is only a short run deviation.

This paper studies the evolution of U.S. property values across 45 metropolitan areas from 1980-2012. It uses a version of the Gordon dividend discount model, modelling price as expected present value of imputed rents. This allows for a parsimonious specification, using only lagged, property values, interest rates, and rents, which are summary statistics for all sorts of variables commonly used in modeling real estate, to explain property values. .

We use Mean Group (MG) and Pooled Mean Group (PMG) estimation, which allow panel data such as ours to have different adjustment speeds across, in our case, cities, but constrain long run pricing (the “fundamentals”) to look like the Gordon model. This means that we can impose standard asset pricing theory in the long run, while allowing for sluggishness of price adjustment and variation of adjustment speeds across cities. We use this structure to analyze regime shifts over the period and differences in “bubbles” across cities, and we test whether the Gordon model applies consistently in the long run and speed of adjustment to the long run.

We find that long run behavior is consistent with fundamentals, and with sensible magnitudes of coefficients and adjustment coefficients that are remarkably similar across cities, but small. Strong cycles are not unusual, nor are serially correlated residuals from our basic equation and "overshooting". Different cities have quite different levels of "momentum," which are correlated with local supply elasticity. We find that the boom from 2002-2006 was longer than usual and had longer positive (in terms of house prices) deviations from our estimated equations. Closer examination of the residuals reveals that the boom appeared to be on the verge of cooling off around 2002 or 2003, but suddenly started up in a way that is consistent with stimulus from the newly-emerging subprime securitization business.

1. Fundamental Models and Models for Estimation

1.1 Modeling the Fundamentals of House Price Growth

Given an information set, Ω_t, the equilibrium condition for holding property at time t is given by[1]

$$P_t = E(R_t|\Omega_t) + E(P_{t+1} / D_t|\Omega_t) \tag{1}$$

where R_t is the net rental income, in this case imputed services of the property, and D_t is the risk-adjusted discount factor. This says that the value of the property is the dividend (rent) plus discounted sale price at the end of the period. Alternatively, it implies that the return on holding property for one period (dividend plus capital gains) must equal the appropriate risk-adjusted hurdle rate.[2] Because future price depends on future rent, this can be iterated to obtain:

$$P_t = \sum_{i=0}^{\infty} E(R_{t+i} / D_{t+i}^i|\Omega_t) + \lim E(P_{t+1+i} / D_{t+1+i}^i|\Omega_t) \tag{2}$$

We define the "fundamental" value as the price given in the case where the transversality condition — that the second term approaches zero — holds. This gives:

$$P_t = \sum_{i=0}^{\infty} E\left(R_{t+i} / D_{t+i}^i|\Omega_t\right). \tag{3}$$

Divinding through by R_t, we have

---

[1] See Lai and Van Order 2010)

[2] See Dougherty and Van Order (1982) for a derivation in the nonstochastic case.

$$P_t / R_t = \sum_{i=0}^{\infty} E\left((R_{t+i} / R_t) / D_{t+i}^{i} \middle| \Omega_t\right) \qquad (3')$$

which corresponds to a price earnings ratio for equities. In what follows we estimate models of the reciprocal of this, R_t/P_t.

For expression (3') to be applicable to owner-occupied housing, imputed rent must be measurable by some form of market rent. We take this to be the actual market rent of comparable properties, which holds if the equation is applied to an owner who is just indifferent between owning and renting. In that case, the first order conditions for both owners and renters will be the same; and the present value formulation that applies to landlords' valuation will apply to owner-occupiers' valuation. The advantage of this approach is that it saves having to develop a complicated model of housing demand and supply, which is not likely to be stable across regions or time. For instance, Glaeser *et al* (2005) emphasizes the role of inelastic supply in house price growth, especially due to local policy variation. Our rent variable captures this effect without having to estimate supply elasticities across cities and time. We measure momentum via lagged effects of rent divided by price. We expect coefficients of these momentum variables to be greater in cities with less elastic supplies.

1.2 Model for Estimation

Equation (3), or (3'), is potentially quite complicated because of complicated adjustment processes. For instance, we should expect interest rates and future rents to be correlated because, given rents, a rise in interest rates will lower property values. On the other hand, an increase in interest rate will induce less production in the future, and thus higher rents. Indeed, if supply is perfectly elastic in the long run, a rise in interest rates will eventually produce a rise in rents without change in long run price. Hence, there is a complicated adjustment process that can be expected to vary significantly across cites with different geography, regulation and so on. However, theory tells us that we should expect prices ultimately to be the expected present value of rent everywhere.

We consider first a simple model with constant interest rates and a steady growth rate of expected rents. Expression (3') is approximated by the Gordon model used for pricing stocks. That is, if interest rates are constant and rents grow at a constant rate, rearranging the formula in (3) we have:

$$\frac{R_t}{P_t} = \alpha_i i_t - \alpha_\pi \pi_t * + \alpha \equiv \gamma_i i_t + \gamma_r r_t + \alpha \qquad (4)$$

where i is the interest rate, $\pi *$ is the expected rates of growth of rent, $r = (i_t - \pi *)$ is the real rate, and α, α_i, α_π, γ_i, and γ_r are parameters. $\frac{R_t}{P_t} + \pi *$ is the expected nominal return on housing. We expect this to be the determinant of long run house prices relative to rents. The term $(\alpha_i i_t - \alpha_\pi \pi_t * + \alpha)$ is the "cap rate" for housing and $(\gamma_i i + \gamma_r r + \alpha)$ is its representation using real interest rates.

In the usual Gordon model the coefficients of i and $\pi *$(and r) are unity. Here we allow for possible tax and other effects that can change such conditions. For instance, if the focus is on the tax break for financing owner-occupied housing but not taxing capital gains on housing, then

$$\frac{R_t}{P_t} = (1-t) i_t - \pi_t * + \alpha \equiv -ti + r + \alpha \qquad (4')$$

where t is marginal tax rate for the marginal homeowner (marginal in the sense of being indifferent between owning and renting). However, it may be the case that high nominal interest rates provide a cash-flow problem for home buyers (even if real rates are constant); then γ_i could be positive.

In general we expect γ_r to be close to 1, α_i and α_π to be positive and not sure about the sign of γ_i. In our data set we have R and P only in the form of indices. Hence, testing for the magnitude of signs (whether they are close to one) especially for γ_r, requires calibrating assumptions. We develop a model that is forced to look like expression (4) in the long run in every city, but allow long adjustments, which vary across cities.

2.3 Dynamic Heterogeneous Panel Estimation

We decompose the relationship in expression (3') into (lagged) long-run and short-run effects among variables using the Pooled Mean Group (PMG) and Mean Group (MG) estimation models developed in Pesaran, Shin and Smith (1997, 1999). The PMG model is restricted maximum likelihood estimation, based on an autoregressive distributed lag (ARDL) model (Pesaran and Shin (1997). Traditionally economic analysis has focused on long run relationships among the dependent variable and the regressors. PMG estimation allows us to identify long run relationships (equation (4)) and short run dynamics relatively easily; the intercepts that reflect the fixed effect, short run coefficients and error variances are allowed to differ across cities, but long run coefficients are constrained to be the same. MG estimation is different in that the long run coefficients are also allowed to vary across cities.

Our model can be represented by:

$$\Delta \frac{R_{c,t}}{P_{ct}} = \sum_{j=1}^{l} \lambda_{c,j,} \Delta \frac{R_{ct-j}}{P_{ct-j}} + \sum_{j=0}^{q} \sum_{k=1}^{n} \delta^{k}{}_{c,j} x^{k}{}_{c,t-j} + \alpha_c + \varepsilon_{c,t} \tag{5}$$

where $\frac{R_{ct}}{P_{c,t}}$ is property rent to price ratio in city c at time t

α_c captures city specific fixed effects

$x^{k}{}_{c,t-j}$ is the kth of n regressors for city c

$\delta^{k}{}_{c,j}$ is the coefficient of the kth regressor for city c

$\lambda_{c,j}$ are scalars

$\varepsilon_{c,t}$ are the city specific errors

c represents panels or cities, $i = 1,2,\ldots,N$

t represents time in quarters, $t = 1,2,\ldots,T$

j is an indicator of lags;

$j = 0,1,2,\ldots,l$ for lagged dependent variable

$j = 0,1,2,\ldots,q$ lags for regressors

Letting $\rho = \frac{R}{P}$ for estimations for (4),[3] (5) can be written as:

[3] A variation is to apply the rationale of expression (6) to expression (5), that is, the dependent variable in (6) is now $\Theta_{j,t} = \frac{R_{j,t}}{P_{j,t-1}} - \frac{R_{j,t}}{R_{j,t-1}}$. Such a model will be comparable to that used in Lai and Van

$$\Delta\rho_{ct} = \lambda_c \rho_{c,t-1} + \sum_{j=0}^{q}\sum_{k=1}^{n} \delta^k{}_{c,j} \Delta x_{c,t-j} + \alpha_c + \varepsilon_{c,t} \tag{6}$$

which when written in error correction form, yields:

$$\Delta\rho_{ct} = \phi_c \left\{ \rho_{c,t-1} - \sum_{k=1}^{n} \beta_c{}^k x_{c,t}^k \right\} - \sum_{j=0}^{q}\sum_{k=1}^{n} \delta_{c,j}^k \Delta x^k{}_{c,t-j} + \alpha_c + \varepsilon_{c,t} \tag{7}$$

where

$$\phi_c = -(1-\lambda_c) \; , \quad \alpha_c = \frac{\mu_c}{(1-\lambda_c)} , \quad \beta_c = \frac{(\delta_{1,c,j} + \delta_{2,c,j})}{(1-\lambda_c)} .$$

Expression (7) is the MG estimation. It allows us restrict some of the parameters inside the brackets to be zero — so we can get to a long run specification that looks like the Gordon model; and it puts no restrictions on parameters across cities. Among the items inside the bracket in (7) are long run fixed effects for each city. However, they cannot be identified separately from the constant terms (the α_c terms in the equation), so we drop them — to be revisited after the α_c have been estimated.

For PMG we assume homogeneous long run relations; i.e., $\beta_c{}^k = \beta^k$ for all *c's* (i.e. cities). Then:

$$\Delta\rho_{ct} = \phi_c \left\{ \rho_{c,t-1} - \sum_{k=1}^{n} \beta^k x_{c,t}^k \right\} - \sum_{j=0}^{q}\sum_{k=1}^{n} \delta_{c,j}^k \Delta x^k{}_{c,t-j} + \alpha_c + \varepsilon_{ct} \tag{8}$$

The double summation term in (8) can have lagged values of changes in the dependent variable, which measure short run momentum.

We estimate variations of expressions (7) and (8). The key is the separation of the model into a long term part in brackets and a short run part, which goes away over time if the model is not explosive. In addition, if PMG in (8) is preferred to MG in (7), we can infer that all MSAs in concern share the same long run effects. Note the model requires rents and prices to grow at the same rate within each city in the long run[4], but the presence of α_c allows the growth rates to vary across cities in the long

Order (2010). However, because this variable is a stationary process, and therefore cannot be cointegrated with interest rates.

[4] We also tried to relax this condition by adding a linear time trend, common to all cities inside the brackets in (8). Results are similar, and therefore are omitted here.

run, which in turn causes the long run level of *R*/*P* to differ across cities. Long run equilibrium is given by:

$$\left\{\rho_{c,} = \sum_{k=1}^{n} \beta^k x_{c,}^k - \alpha_{c/} / \phi_c\right\} \tag{8'}$$

Our approach is ideal for testing the model because it forces the same pricing equation in the long run, with which we force to look like the dividend discount model, but allows different cities to adjust at different rates and in different ways. We test for variation in adjustment paths across cities and for regime shifts over time and by city type (i.e. bubbles versus non-bubbles).

Before testing for the existence of a long run relationship, however, we first need to check if the series are stationary. If some or all the rental income, housing prices, and interest rates are non-stationary, and are integrated of the same order, we can check for their long run relationship with cointegration tests. Hence, the first step is to test if these series are unit roots.

We perform cointegration analysis tests developed by Westerlund (2007) to confirm the existence of long-run relationships among the series. Specifically, Westerlund (2007) relies on the error correction based cointegration. That is, as in expression (5), when ϕ_i, the error correction parameter, is significantly different from zero, then there is a long run relationship (i.e. cointegration). Formally, H_0: $\phi_i = 0$ and H_1: $\phi_i < 0$. Westerlund (2007) proposes four tests. The first two are "group mean statistics" which state that rejecting the null of no cointegration means that at least one or more cities are cointegrated. The test statistics are

$$G_\tau = \frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\phi}_i}{SE(\hat{\phi}_i)}, \quad \text{and} \quad G_\alpha = \frac{1}{N}\sum_{i=1}^{N}\frac{T\hat{\phi}_i}{\hat{\phi}_i(1)} \tag{9}$$

where N is the number of cities, $SE(\hat{\phi}_i)$ is the usual standard error of $\hat{\phi}_i$, and $\hat{\phi}_i(1)$ is the kernel estimator of $\phi_i(1) = 1 - \sum_{j=1}^{l}\phi_{ij}$. The first expression is the t-ratio while the latter is the coefficient statistics (analogous to the rho-statistics of Phillips and Perron (1988).

The other two are "panel statistics," where a rejection of the null of no cointegration means rejection for the panel as a whole. Formally, they are

$$P_{\tau} = \frac{\hat{\phi}}{SE(\hat{\phi})}, \quad \text{and} \qquad P_{\alpha} = T\hat{\phi}. \tag{10}$$

Again, the first expression is the *t*-ratio while the latter is the coefficient statistics.

Westerlund (2007) shows that these statistics are more accurate than the widely used cointegration test due to Pedroni (2004) when the residuals in expression (5), $\varepsilon_{i,t}$, are moving average series.

Given that long run cointegration exists, we next find the long-run and short-run effects among variables using the PMG model. As a robustness check, we also compare the PMG results with MG estimation. The latter assumes that the long run coefficients of each city can be different, and the estimated long run parameter is the average of long run coefficients of all the individual cities. The Hausman test can be used to check if a common long run coefficient exists (that is, not rejecting the null hypothesis of common coefficients between the MG and PMG means common coefficients should be adopted).

2. Tests on Long Run and Regime Shift

Using the Mean Group (MG) estimation in expression (7) and Pooled Mean Group (PMG) estimation in expression (8), we test variations of the fundamental model (4) to see how prices (relative to rents) vary across cities and time. Our measure of house price is the quarterly house price index released by the Office of Federal Housing Finance Administration (FHFA), which provides a repeat sales house price index for over 100 individual Metropolitan Statistical Areas (MSAs) since 1980. The rent series is the "owner's equivalent rent of primary residence" obtained from the Bureau of Labor Statistics, from which we also acquire the local Consumer Price Indices. Figure 1 depicts the rent index to price index ratio.

We use the 10-year Treasury bonds as a measure of nominal long term risk-free rate. We also used the 10-year Treasury Inflation-Protected Securities (TIPS) (bonds issued by the U.S. Treasury that are indexed to inflation) as a direct measure of real interest rates in some variations of the model. This requires assuming expected rent growth to be the same as expected CPI growth, which from Figure 2, appears to be a reasonable assumption, and using it allows elimination of expected inflation from our cap rate. TIPS data are available only after 1998. We therefore interpolate the series back to 1979 Q4, which is explained in Section 3.2. Since it is also possible that market risk could affect the cap rate, we use the Merrill Lynch 1-year high yield rates minus the 1-year Treasury to generate a yield spread to represent market-wide risk.

There is a total of 45 MSAs that have all data available for the required sample period. Since some cities that are more prone to boom might behave differently from those less prone to boom, we follow Lai and Van Order (2010) in classifying the MSAs into bubble MSAs and non-bubble MSAs.

3.1 Panel Unit Root and Cointegration Tests

If property markets were efficient in the usual sense, house prices relative to rents would resemble random walk series, and therefore be non-stationary. If these series are not integrated of order (1) (i.e. $I(1)$), cointegration tests will fail, and therefore MG and PMG estimations cannot be applied. We perform panel unit root tests for the rent to price ratio for different sample periods to check if unit root exists. Several panel unit root tests are adopted here, such as Harris-Tzavalis (1999) test, Breitung test due to Breitung (2000) and Breitung and Das (2005), test due to Hadri (2000), and the IPS, and Fisher-type due to Im, Pesaran and Shin (2003), and Choi (2001) respectively, which are suitable for unbalanced data (not all the MSAs time series have the same length). Except for the test due to Hadri (2000), all tests have the null hypotheses as existence of unit root, and alternative hypotheses as at least one panel stationary. The null hypothesis of Hadri (2000) is that all panels are stationary, while the alternative is to have some panels containing unit root.

Table 1 shows that the rent-price ratio is non-stationary in all the tests, while all the differenced series are stationary, whether de-meaned or not. We also test for stationarity for the interest rate series using Phillips-Perron unit root test and the Augmented Dickey Fuller test, which also show that interest rates are in general non-stationary, or vaguely stationary, while their differenced series are stationary.

We then verify that there is a long run relationship with the Westerlund (2007) panel cointegration test. Even though we attempt to let the tests distinguish bubble MSAs from non-bubble MSAs, considering the possibility of differences in relationships with bubble versus non-bubble MSAs which might cancel each other, we perform cointegration tests separately with non-bubble MSAs and bubble MSAs. Results are shown in Table 2. While the results are not very strong throughout, there is indeed pair-wise cointegration between the rent to price ratio and the various interest rates, particularly with the 10-year Treasury rates, and the 10-year TIPs. The bubble MSAs apparently have stronger cointegration with interest rates than the non-bubble MSAs. For instance, while the overall and the non-bubble MSAs rent to price ratio does not seem to be cointegrated with the high yield rate, there is strong cointegration in the Bubble MSAs. This is a strong hint that bubble MSAs might be riskier than their counterparts, and they might be driven by different forces. Taken as a whole, the tests suggest that the study of long term relationships is feasible.

3.2 Model Estimation and Tests on Regime Shift

We run MG and PMG estimation with variations of expressions (7) and (8), taking account of various lags of the short term variables. We tried 1-, 2-, and 4-lag models and find that the 4-lag (four quarters) models in general are more stable across sub-periods and tests than the other two. We use variations on lagged *R/P* and real and nominal interest rates as short run factors.

It should be noted that an important feature of MG and PMG is that the models allow for regime shifts within the model because there are both long run and short run components, the latter of which should reflect intertemporal differences while avoiding problem of insufficient data series length. Therefore, testing for subperiods to account for possible regime shifts is not essential.[5]

[5] We have actually run the tests for various subperiods including 1980-1990, 1991-1998, 1999-2006, and 2007-2013. Tests results are roughly similar although not very stable possibly because of loss of too many degrees of freedom for such short sample periods. More importantly, while MG estimations

PMG is chosen over MG if a small Hausman test statistics is coupled with a corresponding large *p*-value. That is, the null hypothesis that there is a common long run effect is not rejected, and by having PMG preferred to MG, the long run coefficients are common to all MSAs. Otherwise, MG estimation is better, and the reported long run coefficients are the means of the long run coefficients for individual MSAs.

The residuals from the regressions can provide insights into the bubble formation of the MSAs. For instance, a sum of coefficients of the autoregressive tests of the residuals close to 1 indicates an explosive bubble. Also, a change in the variance of the residuals from these autoregressive tests indicates shift in risk patterns. We therefore run panel autoregressive tests on the residuals from both the MG and PMG results using 2, 4, and 8 lags. Our models do not need very long lags because long run phenomena are reflected in the gradual adjustment to long run effects in the MG and PMG estimations.

We extract the residuals from these panel autoregressive tests to find the variance of the residuals in sub-periods of 1990-1999, 2000-2006, and 2007-2013, which roughly represents pre-bubble, bubble, and post-bubble periods. We then use the Goldfeld-Quandt test to check whether the variances across periods are statistically significantly different, and hence existence of regime shifts. For instance, a variance in the bubble period higher than that of the pre-bubble period indicates increased risk during the bubble period. We also conjecture that volatility of these MSAs after the bubble burst will return to the risk nature in the pre-bubble period. We test this with all MSAs, bubble MSAs, and non-bubble MSAs.

4. Test Results

We use several combinations of long run and short run variables in the MG and PMG tests. The idea is to have different rates to represent the right-hand side of expression (4) because, as seen from Figure 2A, while the few selected interest rate series exhibit

---

are valid, PMG estimations fail in several cases. Even for MG estimations, some variables do not have reasonable coefficients.

similar movements, their differences might generate different effects in our model. Reckoning that the best real rates are actually the TIPs, which however starts only in 1998, we interpolate the data first by regressing TIPS on 10-year Treasury rates and inflation with the data from 1998-2013 and then obtaining the estimated 10-year TIPS series from 1980-1997 based on the regression thus generated. The interpolated TIPS and the actual TIPS are shown in Figure 2B.

For long run variables, we have included combinations of nominal 10-year Treasuries and/or the real interest rates represented by 10-year Treasury Rates minus rent growth, Merryll-Lynch 1-year high-yield spread, and/or TIPS. Short run variables include the lagged rent to price ratio to capture the momentum, and the lagged long run variables to capture the short run effects. In some cases, we also added high yield rates minus rent growth together with the 10-year Treasury rates minus rent growth in the short run variables in order to capture the effects from risky versus riskless rates. We also tried variations of variables such as adding a trend variable to the long-run. Results are similar and therefore selected tests results are discussed here.

All the tests generate consistent explanatory powers from the variables used, with signs as expected. For instance, the error correction coefficients are all negative and significant; and the short run momentum is strong. However, for some of the variations, PMG estimation does not outperform MG estimation (for example, see Appendix B where the Hausman test statistics are significant). In the following, we show the Models that generate the most explanatory power and most number of variables in terms of the rationale of the model as well as the likelihood and consistencies in both MG and PMG estimations. We then choose one of the test results to illustrate the panel autoregression tests for the residuals from the estimations, as well as the variances of the residuals from these AR error equations across subperiods and bubble versus non-bubble MSAs.

4.1 Calibration

As noted above, because rent and price are both indexes we cannot judge magnitudes of coefficients with some sort of calibration assumption. Note from Figure 1 that rent divided by price was close to one for a long period before the boom and bust. It is the case that S & P 500 price-earnings ratio, which is comparable to the inverse of our rent to price ratio, was around 20 for much of this period. Note also that our interest rate numbers are in units like 4.0 rather than .04. Hence, if we take 20 as a comparable and 5% as the earnings to price ratio, then multiplying the coefficients of interest rates on the right hand side by about 5 will calibrate them. Then if our coefficient is .15, we multiply it by 5 to get .75, which in this case is consistent with the impact being close to, but a bit less than, one. This will be discussed below.

4.2 Model Estimation with Nominal Rates —Model A

As our first set of tests, the rationale of Model A is to take on combinations of 10-year TIPS, 10-year Treasuries minus rent growth, and the 10-year Treasuries as long term variables, while lagged values of changes in the ratio of rent to price index, yield-spread, 10-year Treasuries, and 10-year Treasuries minus rent growth are the short term variables. The latter variable choices are mostly to include the same variables as in the long-term variables, plus the lagged values of the dependent variable, the ratio of rent to price index, which would capture the momentum effects.

The MG and PMG results are depicted in Panel 3A of Table 3. Where coefficients are allowed to vary across cities, the table presents averages of the coefficients. In general, the results across the four models are similar. All but one (Model A2) show that PMG performs better than MG (low Hausman Tests coupled with the *p*-values of 10% or higher); that is, the long run variables share the same effect across all MSAs. Even if the test failed to show that PMG is better than MG in Model A2, at least the Model is successful in showing that there is long-run versus short-run relationships in the variables. We shall focus on the PMG results in the following analysis.

The 10-year TIPS dominate the long run effect. The 10-year Treasury minus rent growth, taken to be a way to generate real rates, is not significant, nor is the 10-Year Treasuries, except for Model A3. The error correction coefficients are almost identical throughout. An error correction coefficient of about -0.027 from the quarterly data translates into about 10.8% per annum, meaning that the deviation from the long run is corrected at a rate of 10% each year. The momentum (from the 4 lagged rent-to-price ratios) is strong and significant, with an average sum of coefficients of 0.4. The yield-spread does not show strong short-term effect. What is also interesting is that when both the 10-year Treasuries and the 10-year Treasuries minus rent growth enters the short-run separately, neither is especially significant. But when both are added together, both become significant for all lagged values, but with opposite signs and approximately equal in magnitude. In these cases the long run 10-Year TIPS have lesser effect. Hence, they serve to make up the long effect through short-run influence.

The constant term, which is about 0.027 across the four models, shows the average of the fixed effects of all the MSAs. It is this term that supports our notion that all cities converge to the price given by the Gordon model, but would have sluggishness in price adjustments. The different values of this constant term for different MSAs identify which cities are “growth stocks” relative to others.

Panel 3B summarizes the averages, maxima and minima of the sums of the coefficients of the short run variables, in Model A1, between the bubble and non-bubble MSAs. Sums of coefficients for individual MSAs are listed in Appendix A, Panel A for non-bubble MSAs and Panel B for bubble MSAs. On average, the error correction coefficients of the bubble MSAs (-0.0327) are more negative than those of the non-bubble MSAs (-0.0229). This means that there is bigger correction back to the long run in bubble MSAs.

The average of momentum for non-bubble MSAs is 0.3495 while that of bubble MSAs is 0.5967. In other words, consistent as what has happened, the bubble MSAs have stronger momentum than the non-bubble MSAs. In fact, the difference between the maxima and minima of these sums of coefficients show a wider range for bubble MSAs than non-bubble counterparts. Similarly, the average and the range (maximum – minimum) of the constant term for bubble MSAs are bigger than those of the non-bubble MSAs. This shows that the bubble MSAs are really “growth stocks” that can generate “abnormal returns” on top of the theoretical Gordon model. Yet the momentum of less than 1 means that it is not explosive, even for bubble MSAs.

From the calibration discussion above we get a loose approximation to the magnitude of the effects of real rates (i.e. the long run 10-yr TIPS in Model A) on *R/P* by multiplying the relevant coefficients inside the brackets by 5. The table suggests, then, that the calibrated coefficients are less than one and mostly around .75, which is consistent with the theory.

Figure 3A shows the residuals and the moving averages (with 4 periods) of the residuals of the PMG results of Model A1. The residuals series oscillates quite a lot, which shows autocorrelation. The moving average, which smoothens the series, shows more clearly a regime shift in the residuals for the period of 1999-2006 (the bubble period) and then another shift thereafter (the crisis and post-crisis period). Figure 3B shows the residuals of Model A1 run on bubble and non-bubble MSAs separately. It is obvious that bubble MSAs generate more volatile residuals. However, those of the non-bubble MSAs during the bubble period are smaller than other periods, but fluctuate to a bigger extent in the post-bubble period. This is another trace of possible regime shift. We note that for the long boom from 1999-2006 there is an apparent pause around 2003, which suggests two parts to the boom — the second one roughly coincides with the rise of the subprime securitization market.

We ran residual tests on all models; here we explain the effects of the variables focusing on the results from Model A1. Panel 3C depicts the Autoregression equations of the residuals from the PMG estimation. We test AR(2), AR(4), and AR(8) equations for three sub-periods, 1990-1998, 1999-2006, and 2007-2013, for all MSAs, and for bubble and non-bubble MSAs separately. In general, the AR(2) equation shows the worst explanatory power. We sum all the coefficients of the lagged error term in Panel 3D. Also listed in the same Panel are the variances of the residuals from these autoregressive equations.

The results for all MSAs show that the "bubbles" were not explosive because the sums of the coefficients are all less than 1 (see Panel 3D). The sums of the coefficients of the lagged residuals increased in the bubble period (1999-2006), and then decreased after the Financial Crisis. However, when further testing the equations by separating bubble MSAs from non-bubble MSAs, the residual autoregression in the bubble MSAs dropped significantly after the bubble burst. This decrease is even more severe and becomes negative, a level lower than the pre-crisis period, for bubble MSAs. Those of the non-bubble MSAs are higher than the pre-bubble period, and the drop from the crisis period is either minimal (for 2-lag model) or actually increases for the other two models.

The variances of the residuals from the autoregression equations of bubble MSAs are bigger than those at of non-bubble MSAs, indicating that the former experienced bigger price fluctuations than the latter. Means of the standard deviations of the residuals across cities reveal that bubble MSAs are mostly more volatile than the non-bubble MSAs, as expected. Also, the variances of the residuals are smaller in the bubble period (1999-2006) than in the pre-bubble period. They then increase after the bubble burst. A possible explanation is that volatility tended to be lower during the bubble period when people unanimously anticipated the housing price to rise, but prices fluctuated more after the Crisis.

The question is how different are these variances. Panel 3E shows the comparison with the Goldfeld-Quandt test. Across the Table is the comparison with different autoregression equation for the different sets of MSAs. The different *AR* equations mostly do not generate different variances except during the bubble period. Of more importance is the second half of the Table, which compares the variances across the different periods. In “All MSAs” case, the variances in the pre-bubble and bubble periods are significantly different, and for bubble versus post-bubble period. Exception is the variance of bubble MSAs during and after the bubble periods. In general, there is a significant regime shift in the structure of the housing price fundamentals, particularly before and after the bubbles are formed.

4.2 Explanatory Power of the Model

In the previous subsection, we analyzed the PMG model that provides the long run and short run analysis, and how an additional residual autoregression model can further capture the movement of the rent to price ratio. The issue is how much explanatory power each of these components can offer. To answer this, we compare the sum of squares of the estimated values from the PMG model to the actual rent to price ratios, similar to finding coefficients of determination (R^2). Furthermore, to separate the short run from the long run effects, we borrow the concept of coefficients of partial determination (Partial r^2). Note that the total explanatory power is not the sum of the two coefficients of partial determination (see Borcard (2002)).

Figure 4 shows the fraction of the actual rent to price ratio explained by the PMG model and the residual autoregression model. Figure 5 shows the coefficients of partial determination of long run and short run components. Both figures consist of two panels, Panel A for Non-bubble MSAs and Panel B for Bubble MSAs.

It is obvious that the PMG Model performs better in explaining the bubble MSAs. In fact, the mean explanatory power shown in Figure 4 is 30.34% for non-bubble MSAs and 41.98% for bubble MSAs. Furthermore, Figure 5 shows that the fundamentals as represented in the long run component do not explain much of the rent to price ratio. Instead, it is the short run including momentum that does most of the job. The mean long run and short run coefficients of partial determination for bubble MSAs, shown in Table 4, are 19.54% and 64.72%; while that of the non-bubble MSAs are 8.07% and 51.45%. Besides, the short run component explains more of the bubble MSAs than the non-bubble counterparts, as seen from higher partial r^2 in most of the former MSAs. Finally, Figure 6 shows residuals from the residuals, which suggests something unusual about the 2003-2006 bubble period.

4.3 Model Estimation with Other Rates — Models B and C

Table 5 lists the PMG and MG results for two other sets of Models for comparison. Model B uses only the 10-year TIPS, while Model C has similar long run variables as Model A except replacing the 10-year TIPS with yield spread.

Model B echoes the results of Model A in that the 10-year TIPS exerts long-run effects to the housing rent-to-price dependent variable. Other analyses are similar to the four variations of Model A. On the other hand, all the long run variables in Model C are significant, showing that these variables have to serve the effects that the 10-year TIPS have been influencing, but absent in here. In sum, the different variations of the different models that we have tested all show strong effects from the long-run interest rate variables, with the TIPS fitting the explanation most, and therefore supports the fundamental model.

4.4 Robustness Check of Classification of Non-bubble versus Bubble MSAs

We have so far classified the MSAs into bubble MSAs and non-bubble MSAs based on their price growth rate. Saiz (2010) proposes that supply elasticities can be estimated with the amount of developable land, and that such elasticities are inversely related to housing prices. Here we use the elasticities to test for their correlation with our momentum measure by pairing our MSAs with those in Saiz (2010) to see if there is a negative relationship between our sums of the coefficients of lagged changes in the short-term rent-to-price ratios, which defines the momentum, and his supply elasticities.

The first two panels of Figure 7 plot the momentum of Model A1 and the supply elasticities in Saiz (2010) for non-bubble MSAs and bubble MSAs. It is obvious that the supply elasticities of the non-bubble MSAs are higher than those of bubble MSAs on average. We run an Ordinary Least Squares regression to check their correlations. Panel 7C shows the regression results while Panel 7D plots the actual observations with the regression functions. In all cases – all MSAs, non-bubble MSAs, bubble MSAs, it is the constant that dominates the regressions. Supply elasticity is significant only in the test for all MSAs. The one for Bubble MSAs is even positive. However, it being close to zero in value is far from significant, and therefore can be ignored. Hence, in general, we can conclude that our classification of MSAs is reasonable because the regression result for all MSAs does show that MSAs with higher elasticities (presumably the non-bubble MSAs) have lower prices. Besides, the separate regressions for the bubble MSAs generate a different relationship from the non-bubble MSAs.

5. Comments and Conclusions

We used 34 years of quarterly data from Q1 1980 to Q4 2013 that cover the recent boom and bust period of the US housing markets to estimate if there is a long run relationship between rental rates and interest rates and housing prices. Using Pooled Mean Group estimation (PMG) and Mean Group estimation (MG), we show how housing rent to price ratio adjusts back to the long run fundamental with variations of three different sets of fundamental models. Our results support the argument that housing prices slowly revert to the fundamental model, but the bubble from 1998-2006 helped deviate housing prices farther away from the fundamentals than usual.

The long run fundamentals look very similar across cities and the adjustment speeds are quite similar. However, there are important short run differences in adjustment that are correlated with local supply elasticity. The period after the Great Recession does not see the regime shifting back to that before the formation of the bubbles.

Examining the residuals suggests important points about the recent bubble. Our model has both long run fundamentals to which cities adjust in the long run and short run dynamics that are fundamental to the adjustment processes in the cities. The residuals represent deviations from both of these. The 1999-2006 bubble began like other cyclical changes and in both figures looked like it was coming to an end around 2003. This period does not look like a bubble. However there was a second surge of negative residuals that were what made this cycle worse than others. Our data cannot tell why, but the timing is consistent with the story that it was caused by the sudden surge in subprime (and Alt-A) securitization.

References

Black, A., P. Fraser and M. Hoesli. (2006). House Prices, Fundamental and Bubbles. *Journal of Business Finance and Accounting* 33:1535-1555.

Borcard, Daniel (2002). Multiple and partial regression and correlation Partial *r*2, contribution and fraction [a]. http://biol09.biol.umontreal.ca/borcardd/partialr2.pdf.

Breitung, Jorg (1999). The Local Power of Some Unit root Tests for Panel Data, Discussion Papers, Interdisciplinary Research Project 373: Quantification and Simulation of Economic Prcesses, No. 1999, 69, http://nbn-resolving.de/urn:nbn:de:kobv:11-10046584

Breitung, Jorg and Samarjit Das (2005). Panel Unit Root Tests Under Cross-sectional Dependence, *Statistica Neerlandica*, 59(4), 414-433.

Choi, In (2001). Unit Root Tests for Panel Data, *Journal of International Money and Finance*, 20, 249-272.

Chan, I.C., S.K. Lee and K.Y. Woo. 2001. Detecting Rational bubbles in the Residential Housing Markets in Hong Kong. *Economic Modelling* 18:61-73.

Coleman, M.IV, M. LaCour-Little and K.D. Vandell. 2008. Subprime Lending and the Housing Bubble: Tail Wags Dog? Social Science Research Network: http://ssrn.com/abstract=1262365.

Chang, Y., A.C. Cutts and R.K. Green. 2005 Did Changing Rents Explain Changing House Prices During the 1990s? George Washington University School of Business Working Paper.Glaseser, Edward, Joseph Gyourko and Raven Saks (2005). Why Have Housing Prices Gone Up? *American Economic Review*, 95(2), 329-333.

Hwang, M., J. Quigley and J. Son. 2006. The Dividend Pricing Model: New Evidence from the Korean Housing Market. *Journal of Real Estate Finance and Economics* 32:205-228.

Im, Kyung So, M. Hashem Pesaran, Yongcheol Shin (2003). Testing for Unit Roots in Heterogeneous Panels, *Journal of Econometrics*, 115, 53-74.

Lai, Rose, N. and Robert Van Order. 2010. Momentum and House Price Growth in The U.S: Anatomy Of A Bubble, *Real Estate Economics*, 38(4), 753-773.

Pesaran, Hashem M., Yongcheol Shin and Ron Smith. 1999. Pooled Mean Group Estimation Of Dynamic Heterogeneous Panels. *Journal of the American Statistical Association*, 94, 621-634.

Pesaran, Hashem M., and Yongcheol Shin. 1999. An Autoregressing Distributed Lag Modelling Approach to Cointegration Analysis. (ed) S Strom, *Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium*, chapter 11. Cambridge University Press, Cambridge.

Pesaran, Hashem M. and Ron Smith. 1995. Long run Relationships from Dynamic Heterogeneous Panels. *Journal of Econometrics*, 68, 79-113.

Saiz, Albert. 2010. The Geographic Determinants of Housing Supply, *Quarterly Journal of Economics*, 125(3), 1253-1296.

Taipalus, K. 2006. A Global House Price Bubble? Evaluation Based on a New Rent-Price Approach. Bank of Finland Discussion Papers. 29.

Wheaton, W. and G. Nechayev. 2008. The 1998-2005 Housing 'Bubble' and the Current 'Correction': What's Different this Time? *Journal of Real Estate Research* 30:1-26.

Figure 1 Ratio of US National Rent Index to Price Index

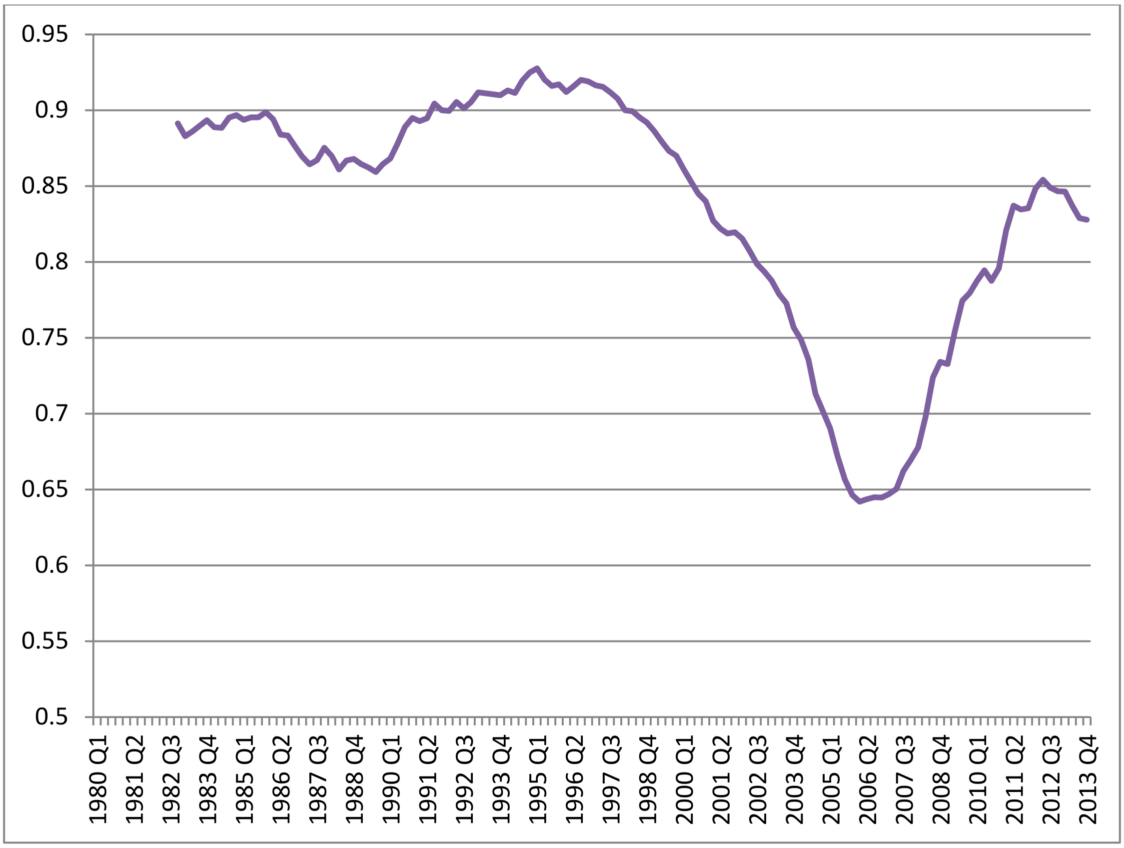
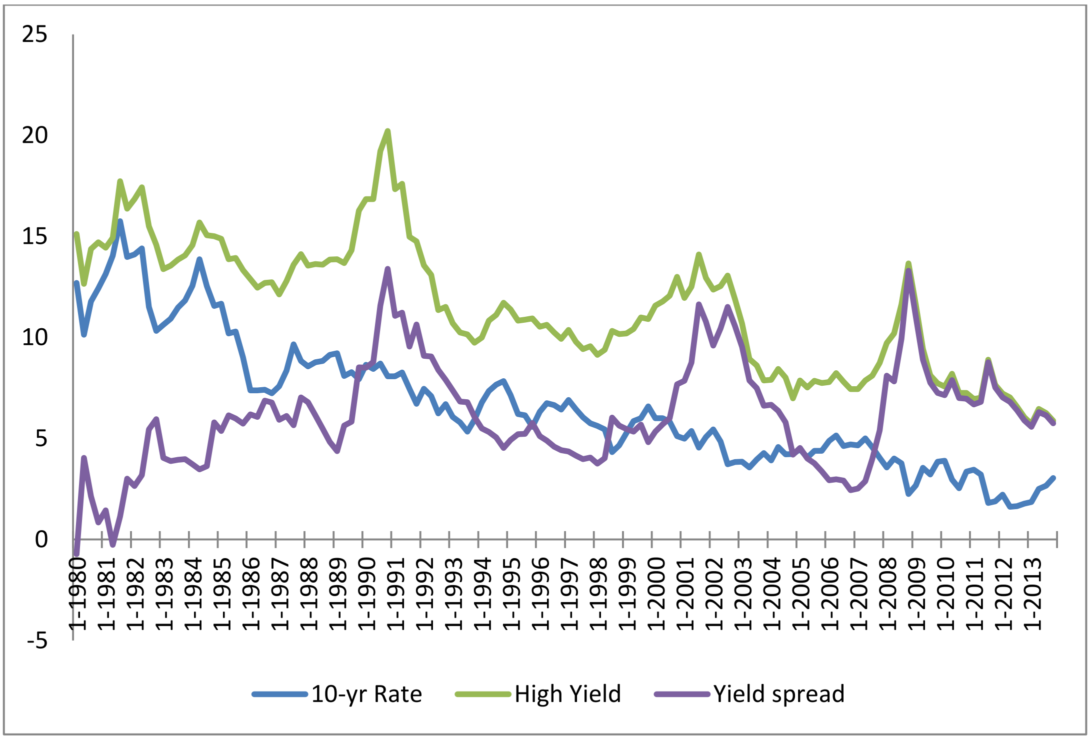


Figure 2 Movements of Various Rates Used for the Tests

Panel 2A Plots of 10-year Treasuries, Merrill Lynch High-yield Rate, and Corresponding Yield Spread



Panel 2B Plots of Actual TIPS and Interpolated TIPS

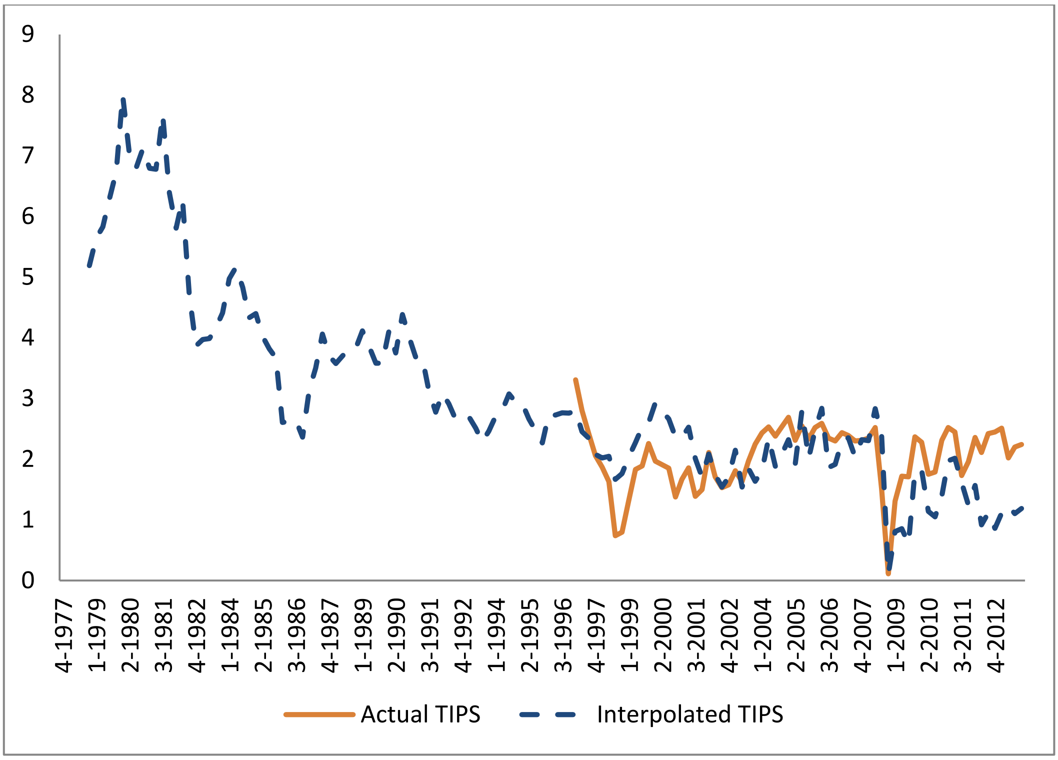
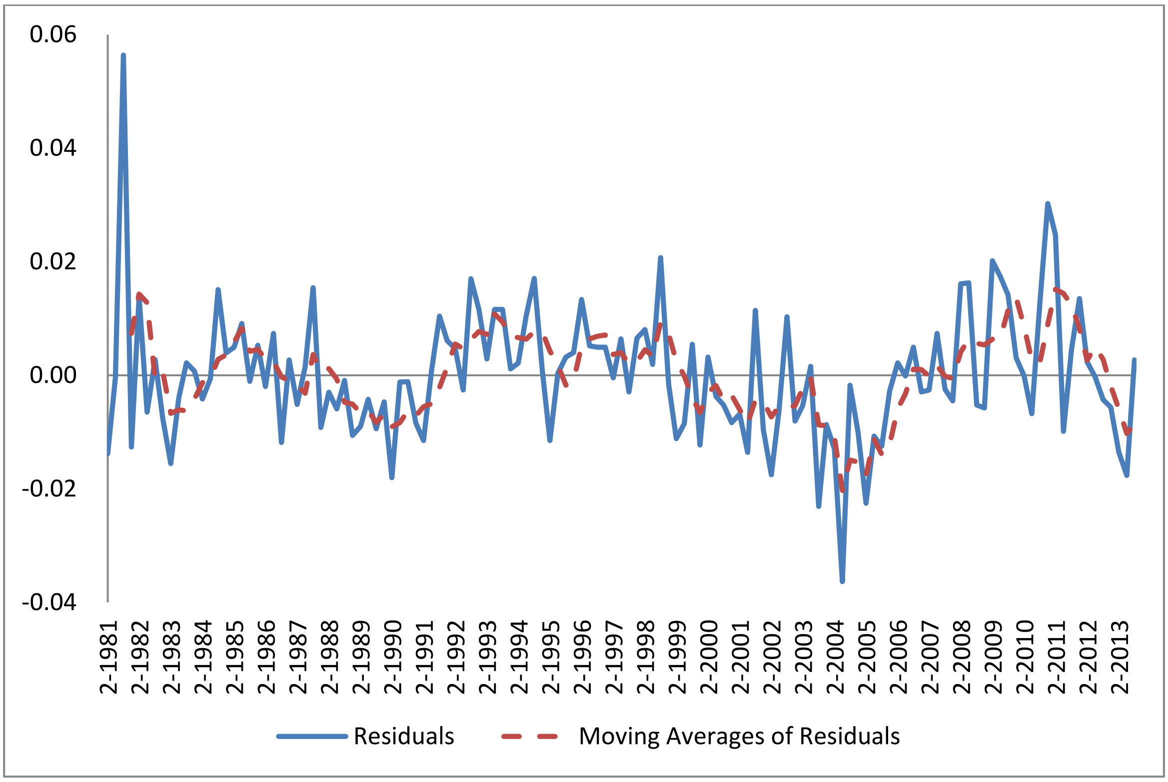


Figure 3 Residuals from Model A1 PMG Estimation

Panel 3A Average and Moving Averages of Residuals



Panel 3B Residuals of Bubble versus Non-Bubble MSAs

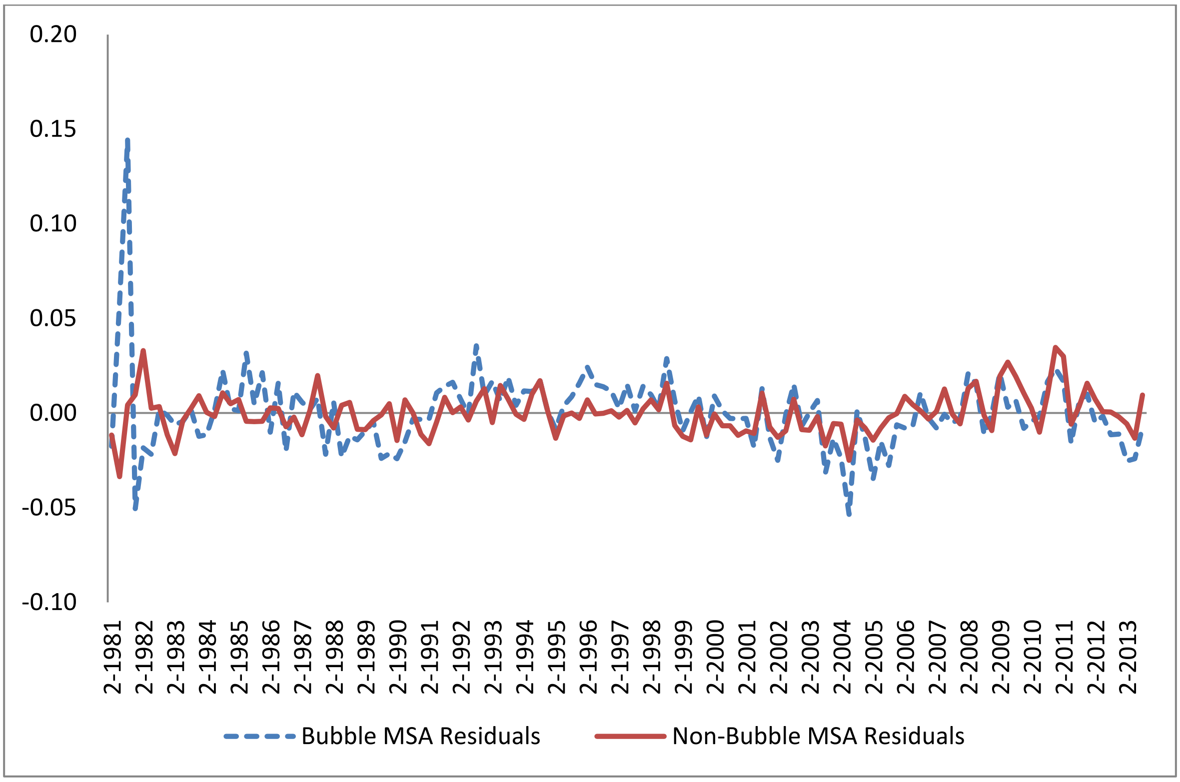
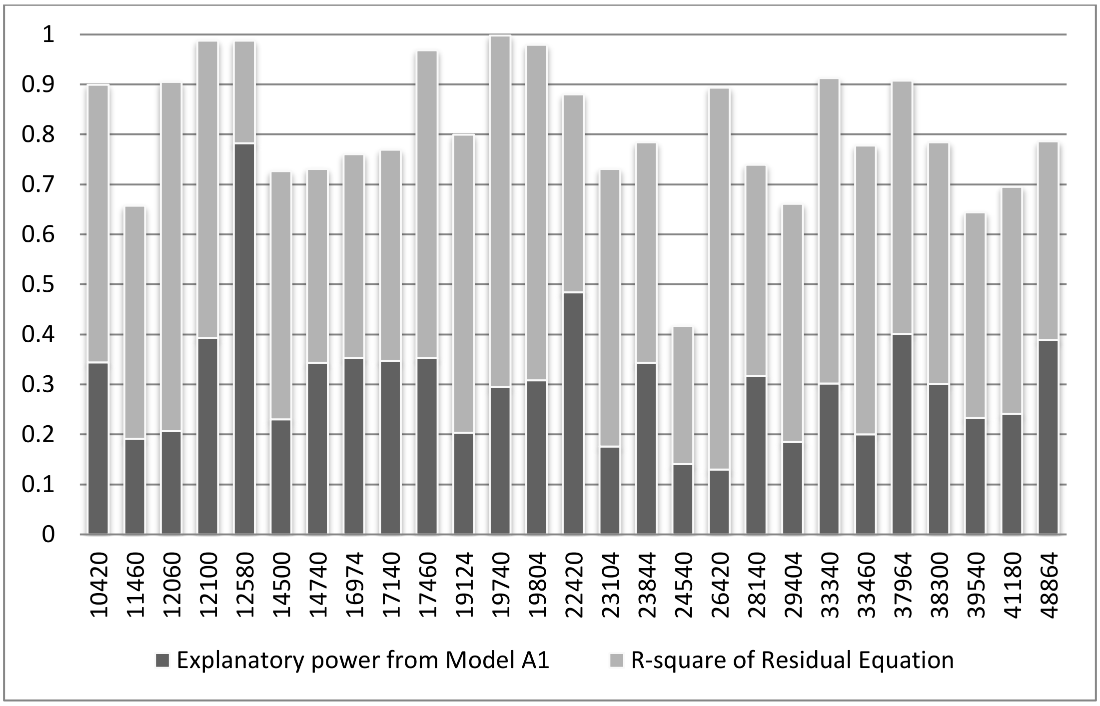


Figure 4 Explanatory Power of PMG and Residual Autoregression of Model A1

Panel 4A Non-bubble MSAs



Panel 4B Bubble MSAs

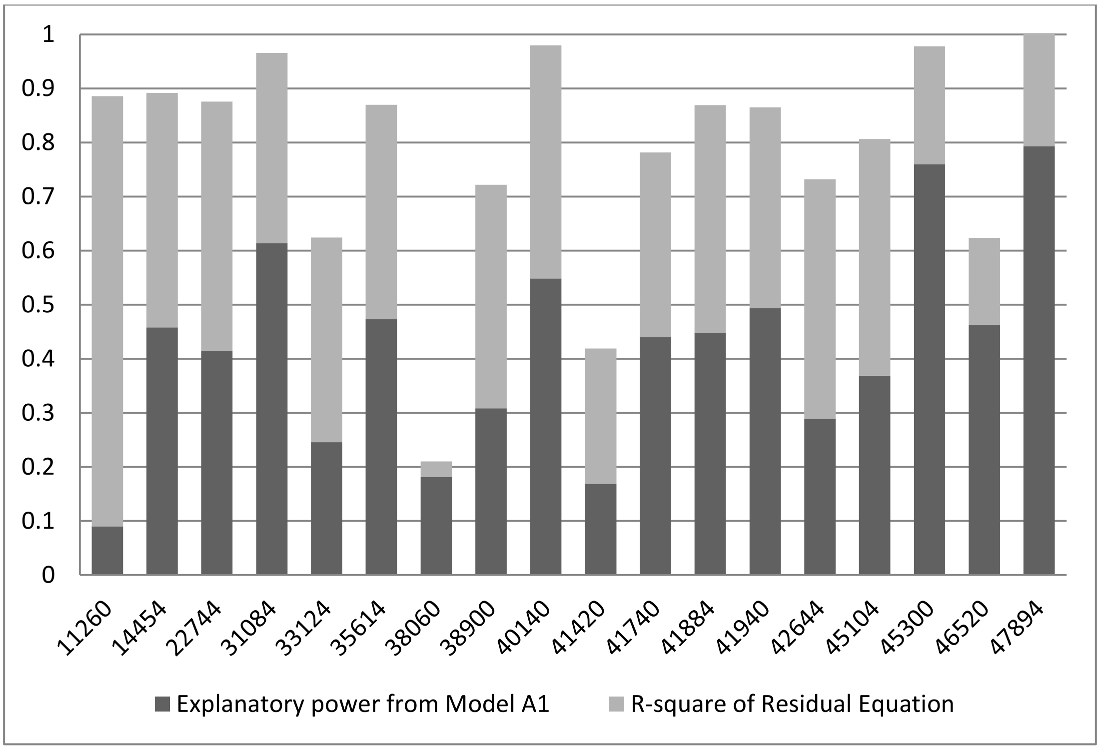
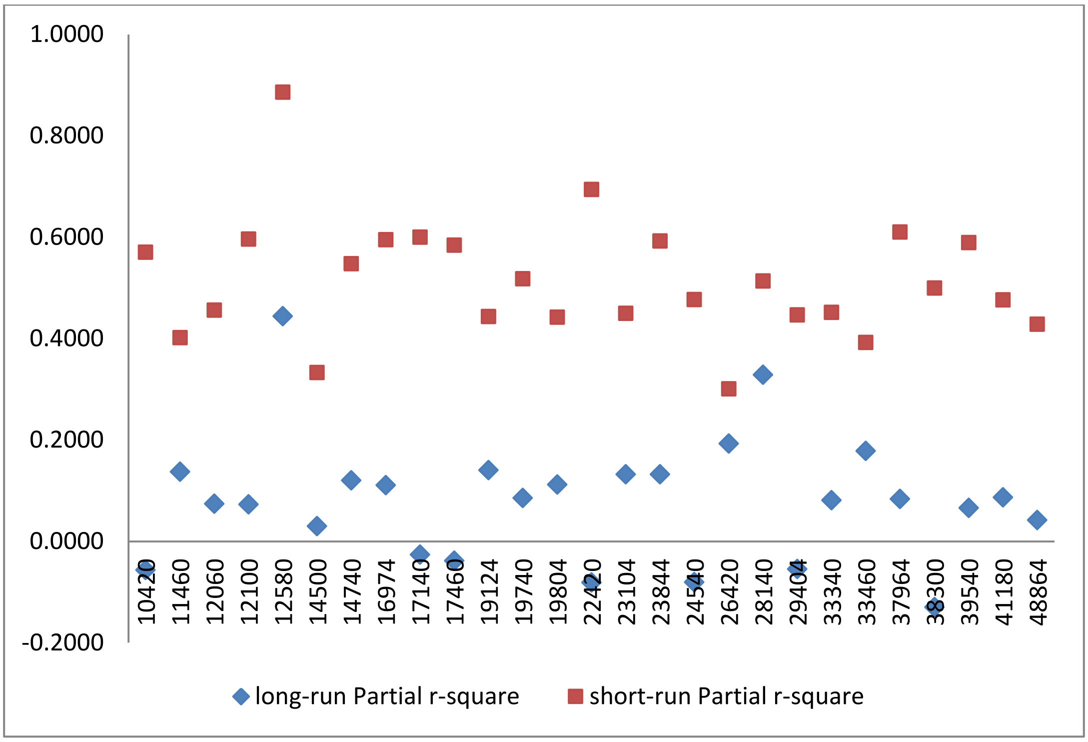


Figure 5 Partial r^2 of Long Run and Short Run Components of Model A1

Panel 5A Non-bubble MSAs



Panel 5B Bubble MSAs

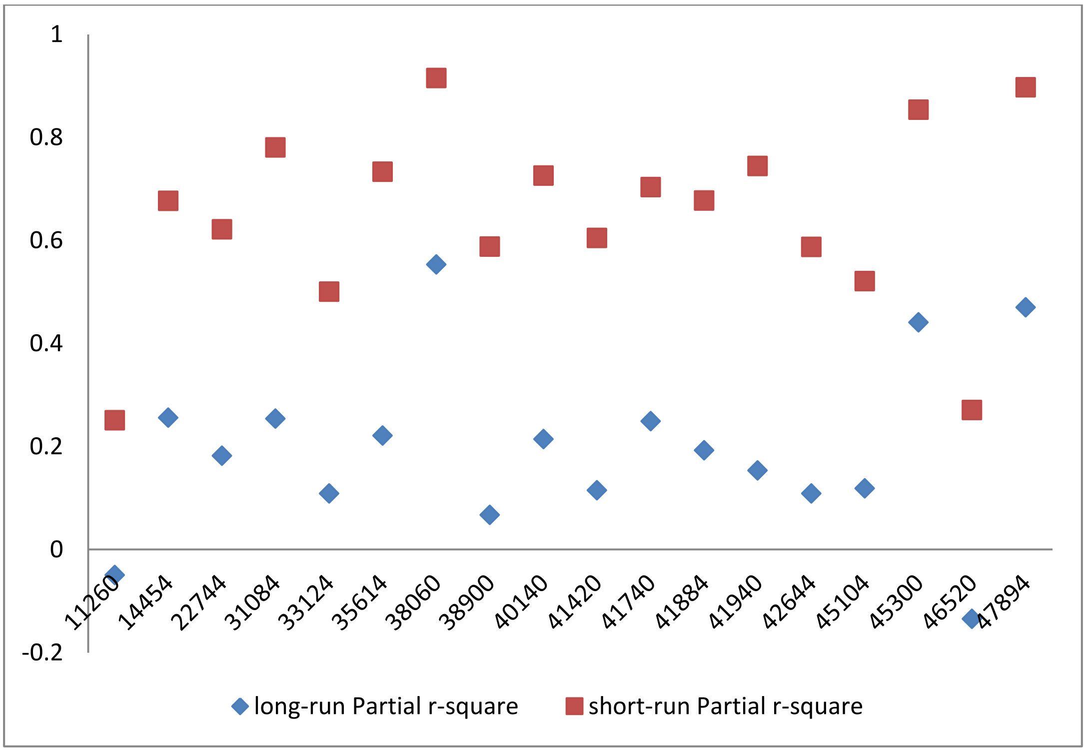
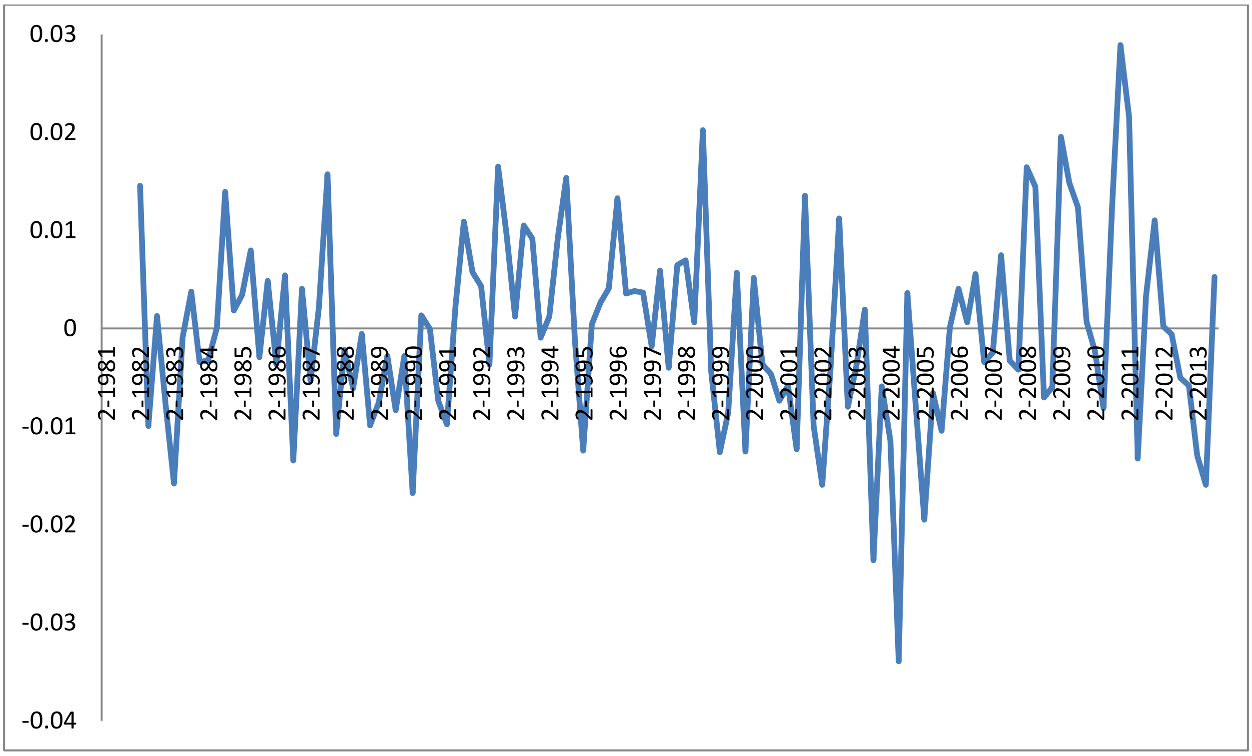


Figure 6 **Residuals of the *AR*(4) Residual Equation from Model A1**

Panel 6A ***All MSAs***



Panel 6B ***Bubble and Non-Bubble MSAs***

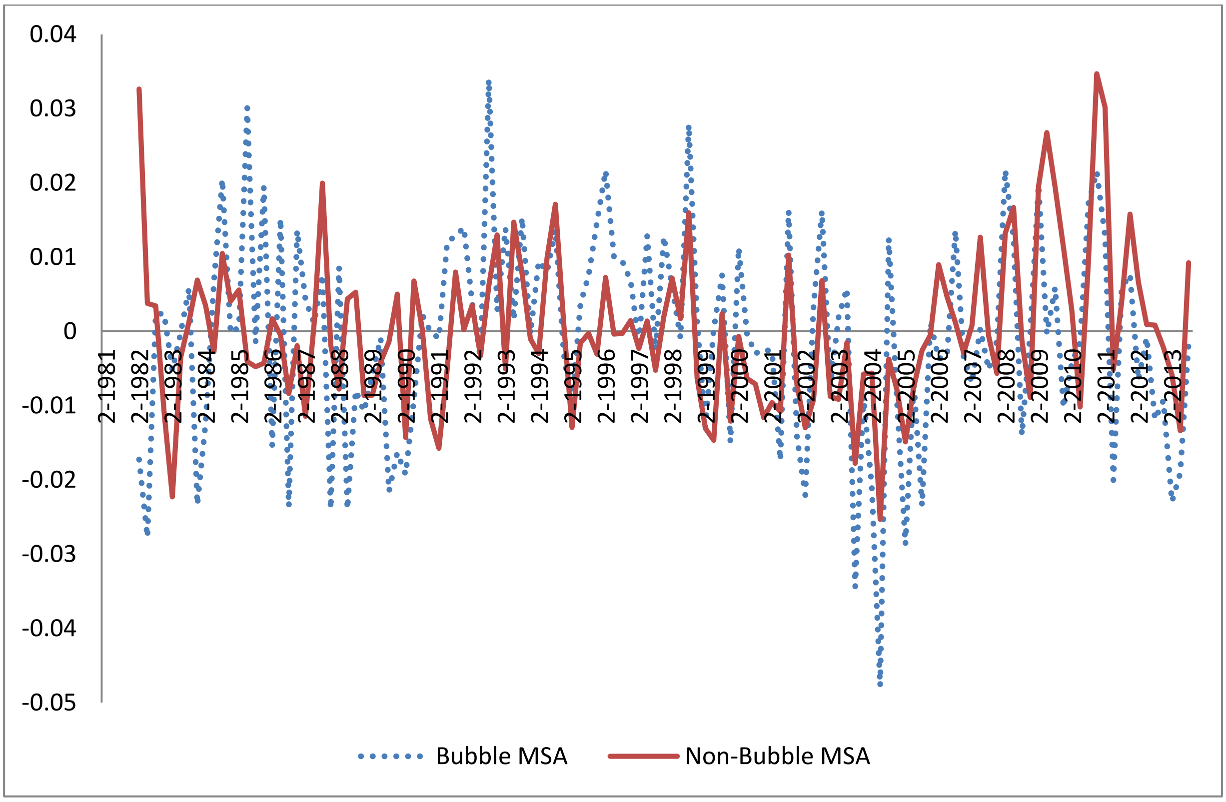
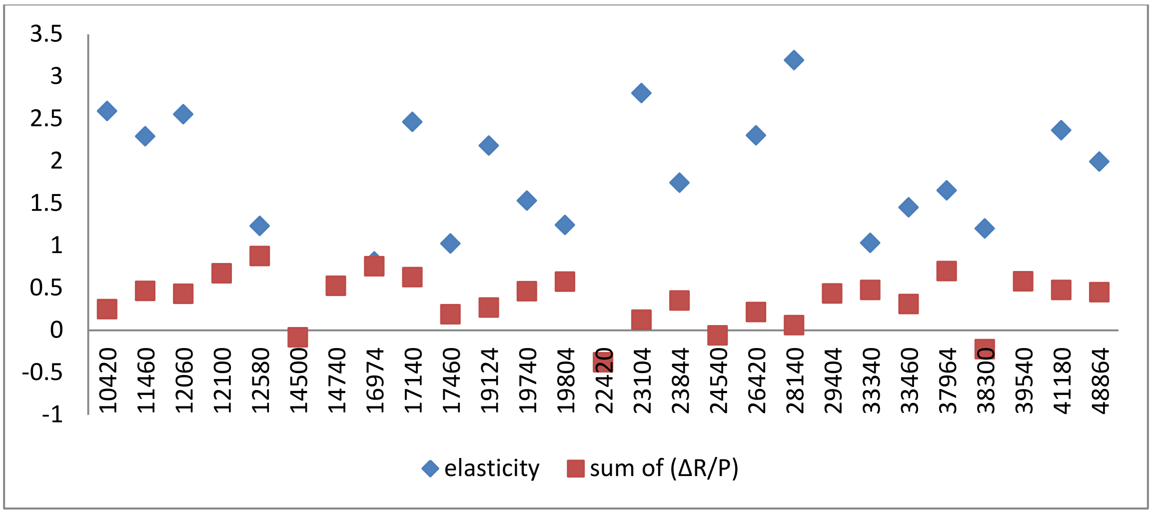
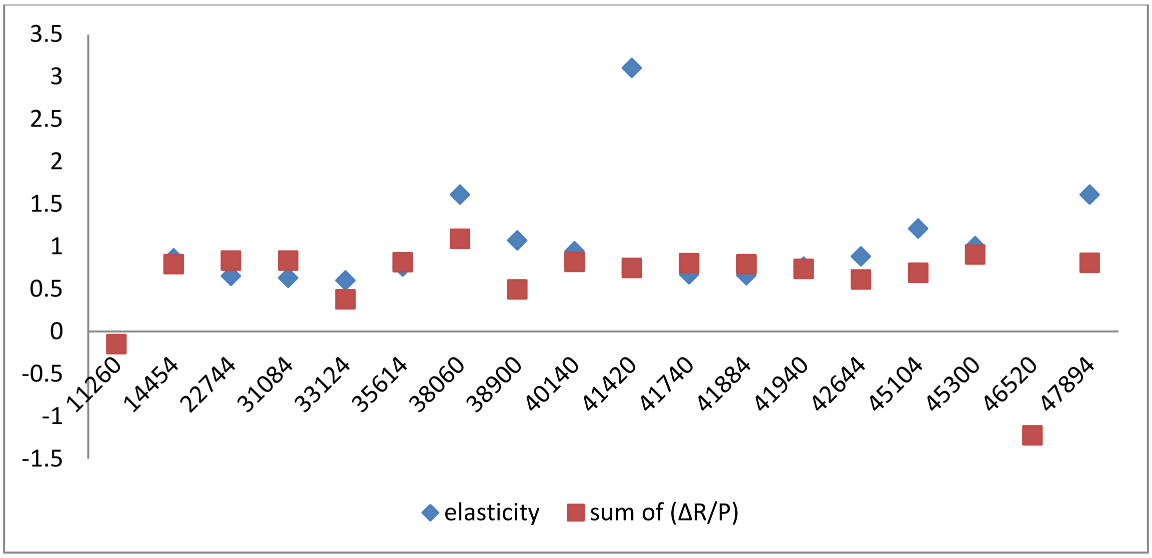


Figure 7 **Relationship between Sum of Changes in Rent-to-Price Ratio and Supply Elasticity (Saiz, 2014)**

Panel 7A ***Relationship for Non-Bubbles MSAs***



Panel 7B ***Relationship for Bubbles MSAs***



Panel 7C ***OLS Regression of Sum of Changes in Rent-to-Price Ratio on Supply Elasticity***

	sum of (ΔR/P)		
	All MSAs	**Non-bubble MSAs**	**Bubble MSAs**
Supply elasticity	-0.1660***	-0.1138	0.0414
	(0.006)	(0.196)	(0.554)
Constant	0.8045***	0.6028***	0.7132***
	(0.000)	(0.002)	0.000
Observations	36	20	16
R-squared	0.199	0.091	0.026
Adjusted R-squared	0.176	0.0405	-0.044

Panel 7D *Plots of the OLS Regression and Actual Observations of Sum of Changes in Rent-to-Price Ratio on Supply Elasticity*

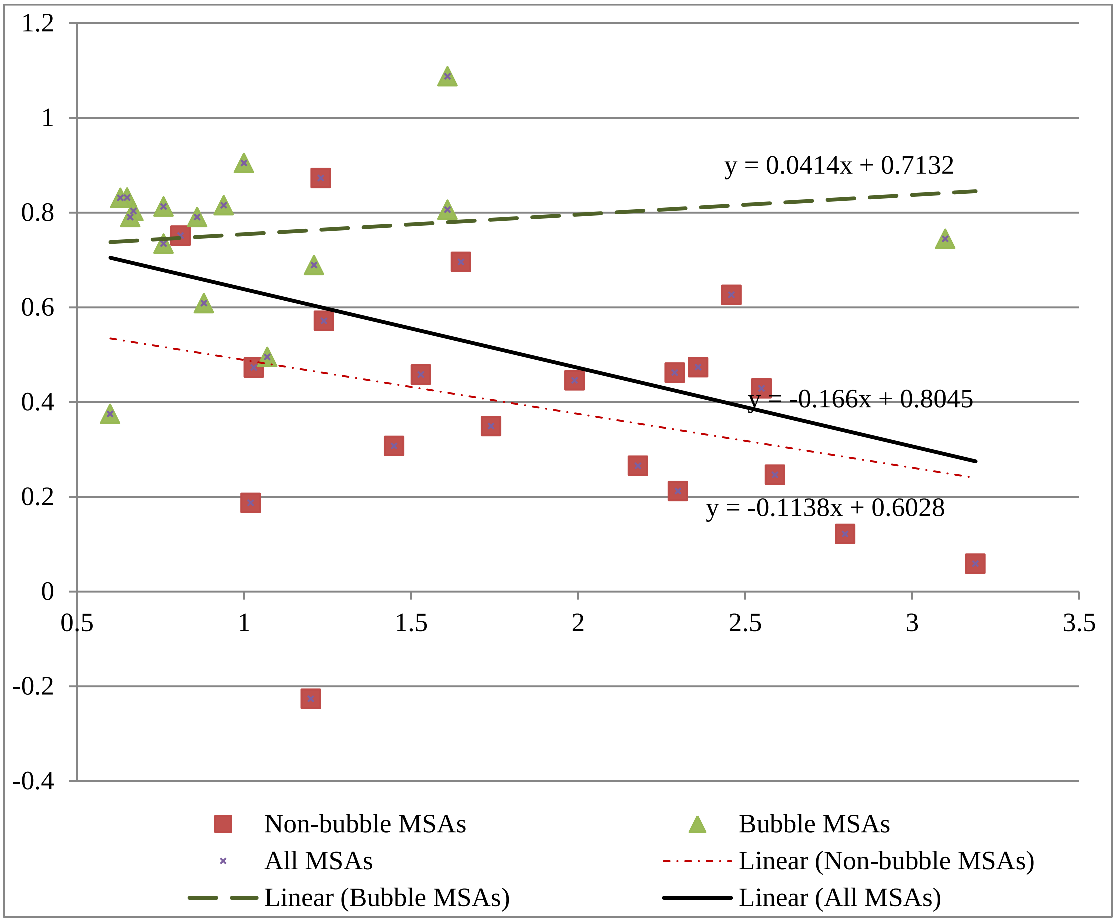


Table 1 Tests for Stationarity

Panel 1A Tests for Non-Stationarity

Tests	Rent to Price Ratio		Differenced Rent to Price Ratio	
	DEMEAN	not-DEMEAN	DEMEAN	not-DEMEAN
Harris–Tzavalis (1999)	0.9788	0.990	0.1752***	0.101***
Breitung (2000); Breitung and Das (2005)	3.071	3.025	-9.6508***	-7.335***
Im–Pesaran–Shin (2003)	5.869	9.583	-27.7329***	-40.242***
Fisher-type (Choi 2001)	-3.815	-4.946	69.5614***	109.792***
Hadri (2000)	39.582***	31.612***	9.7717***	13.660***

Note: "***" denote significance at 1% levels. Except for Hadri LM test, all tests have H_0: All panels contain unit roots. For Hadri LM test, H_0: All panels are stationary. Ha: Some panels contain unit roots.

Panel 1B Unit Root Tests for Various Interest Rates

	On Variables		On Differenced Variables	
Variables	**ADF test**	**Phillips-Perron test**	**ADF test**	**Phillips-Perron test**
10-Year rates	-1.369	-1.337	-12.931***	-12.895***
10-Year TIP rates	-4.027***	-4.096***	-7.949***	-8.062***
High Yield rates	-1.414	-1.669	-10.995***	-11.051***
High Yield Spread	-2.809*	-2.963**	-13.117***	-12.970***
Generated 10-Year TIPs	-2.697*	-2.639*	-12.542***	-12.593***

Note: *, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

Table 2 Cointegration Tests Of Rent to Price and Various Rates Based on Westerlund (2007)

	G_τ	G_α	P_τ	P_α
Rent to Price Ratio and 10-year rate				
All MSAs	-1.413***	-1.650	-8.576***	-1.486
Non-bubble MSAs	-1.036	-0.814	-4.539**	-0.713
Bubble MSAs	-2.110***	-3.193	-7.63***	-3.166***
Rent to Price Ratio and 10-year TIPs				
All MSAs	-2.133***	-1.91	-13.643***	-2.112**
Non-bubble MSAs	-1.958***	-1.561	-10.717***	-1.75
Bubble MSAs	-2.456***	-2.555	-8.443***	-2.547**
Rent to Price Ratio and High Yield rate				
All MSAs	-0.880	-1.709	-6.97***	-1.732*
Non-bubble MSAs	-0.266	-0.400	-1.443	-0.329
Bubble MSAs	-2.014***	-4.125***	-7.442***	-4.17***
Rent to Price Ratio and High Yield Spread				
All MSAs	-0.566	-0.575	-4.292	-0.674
Non-bubble MSAs	-0.323	-0.253	-1.675	-0.259
Bubble MSAs	-1.014	-1.17	-3.877**	-1.272
Rent to Price Ratio and Interpolated 10-year TIPs				
All MSAs	-1.868***	-2.262	-11.531***	-2.255***
Non-bubble MSAs	-1.575***	-1.433	-7.293***	-1.348
Bubble MSAs	-2.41***	-3.791	-8.758***	-3.769***

Note: *, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

Table 3 PMG & MG Estimation for Rent to Price Ratio Model A

Panel 3A PMG and MG Estimation Results

	Model A1		**Model A2**		**Model A3**		**Model A4**	
	PMG	**MG**	**PMG**	**MG**	**PMG**	**MG**	**PMG**	**MG**
Long Run Variables								
10-yr TIPs	0.1428***	0.1549***	0.1571***	0.1602***	0.0728***	0.0812***	0.0749***	0.0814***
10Y- RentG							0.5877	1.0854
10-year Treasury	0.0061	0.0077	0.001	0.0061	0.0424***	0.0338***	-0.5439	-1.047
Short Run Variables								
Error Correction	-0.0268***	-0.0348***	-0.0268***	-0.0350***	-0.0268***	-0.0362***	-0.0271***	-0.0423***
$\Delta R/P_{t-1}$	0.0433	0.0387	0.0496	0.0451	0.2005***	0.1924***	0.2001***	0.1809***
$\Delta R/P_{t-2}$	0.1073***	0.1052***	0.0999***	0.0986***	0.0745***	0.0795***	0.0740***	0.0764***
$\Delta R/P_{t-3}$	0.1677***	0.1694***	0.1651***	0.1671***	0.2002***	0.2043***	0.2000***	0.2025***
$\Delta R/P_{t-4}$	0.1301***	0.1317***	0.1270***	0.1292***	0.1538***	0.1605***	0.1525***	0.1548***
ΔYield spread$_t$	0.0021***	0.0019***	0.0017***	0.0015***	0.0019***	0.0017***	0.0019***	0.0019***
ΔYield spread$_{t-1}$	0.0016	0.0015	0.0018	0.0017	0.0011*	0.0010*	0.0011*	0.0013*
ΔYield spread$_{t-2}$	0.0020*	0.0019	0.0020*	0.0019	0.0019***	0.0018**	0.0019***	0.0020***
ΔYield spread$_{t-3}$	0.0021***	0.0019***	0.0020***	0.0019***	0.0011*	0.0011*	0.0011*	0.0012*
ΔYield spread$_{t-4}$	0	-0.0003	-0.0001	-0.0003	0.0001	-0.0001	0.0001	-0.0001
$\Delta 10Y_t$	-0.0018	-0.0025			0.3565***	0.3619***	0.3689***	0.4514***
$\Delta 10Y_{t-1}$	0.0082**	0.0076**			0.2145***	0.2267***	0.2242***	0.3012***
$\Delta 10Y_{t-2}$	-0.0004	-0.0009			0.2646***	0.2748***	0.2718***	0.3268***
$\Delta 10Y_{t-3}$	-0.0026*	-0.0033**			0.1144***	0.1235***	0.1192***	0.1589***
$\Delta 10Y_{t-4}$	-0.0003	-0.0009			0.0379	0.0422	0.0403	0.0582

Table 3A Continued

	Model A1		**Model A2**		**Model A3**		**Model A4**	
	PMG	**MG**	**PMG**	**MG**	**PMG**	**MG**	**PMG**	**MG**
Short Run Variables								
$\Delta 10Y_t$ - $RentG_t$			-0.0033	-0.0040*	-0.3572***	-0.3631***	-0.3697***	-0.4536***
$\Delta 10Y_{t-1}$ - $RentG_{t-1}$			0.0089**	0.0083**	-0.2072***	-0.2199***	-0.2170***	-0.2950***
$\Delta 10Y_{t-2}$ - $RentG_{t-2}$			-0.001	-0.0015	-0.2660***	-0.2765***	-0.2732***	-0.3289***
$\Delta 10Y_{t-3}$ - $RentG_{t-3}$			-0.0023	-0.0030**	-0.1180***	-0.1276***	-0.1229***	-0.1634***
$\Delta 10Y_{t-4}$ - $RentG_{t-4}$			-0.0004	-0.001	-0.038	-0.0429	-0.0404	-0.0595
Constant	0.0273***	0.0347***	0.0271***	0.0348***	0.0268***	0.0366***	0.0272***	0.0433***
Log Likelihood	12707	12757	12702	12753	13859	13922	13859	13981
Hausman Test	4.52		-2023.61		1.08		0.47	
***p*-value**	0.1041		**Invalid**		0.5818		0.9249	

Note : *, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

The variable 10-Y – RentG represent 10-year Treasury rates minus rent growth, yield spread is Merrill Lynch 1-year high yield bond rates minus 1-year Treasury rates.

An insignificant value of Hausman Test indicates PMG is preferred to MG Estimation. "Invalid" *p*-value is because of failure of Hausman Test.

Panel 3B ***Summary Statistics of Sum of Short-run Coefficients for Individual MSAs from PMG Estimation of Model A1 (Individual MSAs in Appendix A)***

	Error Correction		**sum of ($\Delta R/P$)**		**sum of (ΔYield spread)**		**Constant**	
	Non-Bubble MSA	Bubble MSA	Non-Bubble MSA	Bubble MSA	Non-Bubble MSA	Bubble MSA	Non-Bubble MSA	Bubble MSA
Average	-0.0229	-0.0327	0.3495	0.5967	0.0055	0.0057	0.0048	0.0051
Maximum	-0.0105	-0.0119	0.8728	1.0877	0.0156	0.0122	0.0293	0.1443
Minimum	-0.0379	-0.1059	-0.3830	-1.2240	-0.0008	-0.0053	-0.0158	-0.0133

Panel 3C Residual Autoregressive Models from Model A1

	2 Lags			4 Lags			8 Lags		
	1990-1998	**1999-2006**	**2007-2013**	**1990-1998**	**1999-2006**	**2007-2013**	**1990-1998**	**1999-2006**	**2007-2013**
All MSAs									
$\varepsilon_{i,t\text{-}1}$	0.0315	0.1930***	0.0645**	-0.0301	0.0815***	0.0640**	-0.0346	0.0944***	0.0624**
$\varepsilon_{i,t\text{-}2}$	0.2819***	0.2590***	-0.0058	0.2556***	0.2002***	-0.0109	0.2509***	0.1979***	-0.0332
$\varepsilon_{i,t\text{-}3}$				0.1639***	0.3992***	0.0913***	0.1458***	0.3658***	0.1011***
$\varepsilon_{i,t\text{-}4}$				0.0594**	-0.0206	0.0324	0.0342	-0.0131	0.0209
$\varepsilon_{i,t\text{-}5}$							0.0001	-0.0547*	0.1927***
$\varepsilon_{i,t\text{-}6}$							0.0881***	0.0866***	0.0096
$\varepsilon_{i,t\text{-}7}$							0.0632***	-0.0726**	0.0113
$\varepsilon_{i,t\text{-}8}$							-0.0437*	0.0834***	-0.0581**
Adjusted R^2	0.0819	0.134	0.00254	0.112	0.268	0.0104	0.123	0.28	0.0465
Bubble MSAs									
$\varepsilon_{i,t\text{-}1}$	0.0503	0.1989***	-0.0622	-0.0434	0.0554	-0.0555	-0.0299	0.0676	-0.0479
$\varepsilon_{i,t\text{-}2}$	0.3711***	0.3025***	0.0165	0.3253***	0.2396***	0.0212	0.3523***	0.2645***	0.0178
$\varepsilon_{i,t\text{-}3}$				0.1914***	0.4802***	0.0722	0.1796***	0.4711***	0.0841*
$\varepsilon_{i,t\text{-}4}$				0.0728*	-0.0499	-0.0892**	0.0407	-0.0337	-0.0737*
$\varepsilon_{i,t\text{-}5}$							-0.0844**	-0.1008**	0.1487***
$\varepsilon_{i,t\text{-}6}$							0.0991**	0.0424	0.0005
$\varepsilon_{i,t\text{-}7}$							0.0981**	-0.0653	-0.0621
$\varepsilon_{i,t\text{-}8}$							-0.1014***	0.0758	-0.0909**
Adjusted R^2	0.142	0.168	0.000385	0.18	0.352	0.0103	0.203	0.357	0.0346

(Panel 3C Continued)

	2 Lags			**4 Lags**			**8 Lags**		
	1990-1998	**1999-2006**	**2007-2013**	**1990-1998**	**1999-2006**	**2007-2013**	**1990-1998**	**1999-2006**	**2007-2013**
Non-Bubble MSAs									
$\varepsilon_{i,t\text{-}1}$	-0.0598*	0.1529***	0.3313***	-0.0600*	0.1220***	0.3156***	-0.0709**	0.1329***	0.2495***
$\varepsilon_{i,t\text{-}2}$	0.0569*	0.1123***	-0.1462***	0.0649**	0.0832**	-0.1691***	0.0497	0.0625*	-0.2344***
$\varepsilon_{i,t\text{-}3}$				0.0544*	0.1602***	0.1193***	0.0374	0.1414***	0.1158***
$\varepsilon_{i,t\text{-}4}$				-0.0412	0.0524	0.2118***	-0.0308	0.0455	0.1525***
$\varepsilon_{i,t\text{-}5}$							0.1154***	-0.0339	0.2005***
$\varepsilon_{i,t\text{-}6}$							0.0532*	0.1148***	0.0549
$\varepsilon_{i,t\text{-}7}$							0.0497*	-0.0541	0.1756***
$\varepsilon_{i,t\text{-}8}$							0.0716**	0.1110***	-0.0423
Adjusted R^2	0.00563	0.0412	0.0962	0.0089	0.069	0.163	0.0317	0.0922	0.218

Note : *, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

Panel 3D   Sum of Coefficients and Variances of Residual Autoregression Estimation in Panel B

	2 Lags			**4 Lags**			**8 Lags**		
	1990-1998	**1999-2006**	**2007-2013**	**1990-1998**	**1999-2006**	**2007-2013**	**1990-1998**	**1999-2006**	**2007-2013**
All MSAs									
Sum of all Coefficients	0.3134	0.452	0.0587	0.4488	0.6603	0.1768	0.504	0.6877	0.3067
Sum of significant Coefficients	0.2819	0.452	0.0645	0.4789	0.6809	0.1553	0.5043	0.7008	0.2981
Variance of Residuals	0.00083	0.00060	0.00078	0.00080	0.00051	0.00078	0.00079	0.00051	0.00075
Bubble MSAs									
Sum of all Coefficients	0.4214	0.5014	-0.0457	0.5461	0.7253	-0.0513	0.5541	0.7216	-0.0235
Sum of significant Coefficients	0.3711	0.5014	0.0000	0.5895	0.7198	-0.0892	0.5433	0.6348	0.0682
Variance of Residuals	0.00150	0.00115	0.00123	0.00144	0.00090	0.00122	0.00140	0.00091	0.00119
Non-Bubble MSAs									
Sum of all Coefficients	-0.0029	0.2652	0.1851	0.0181	0.4178	0.4776	0.2753	0.5201	0.6721
Sum of significant Coefficients	-0.0029	0.2652	0.1851	0.0593	0.3654	0.4776	0.2190	0.5626	0.6595
Variance of Residuals	0.00040	0.00025	0.00044	0.00040	0.00024	0.00041	0.00039	0.00023	0.00038

Panel 3E ***Goldfeld-Quandt Tests of Variance of Residuals from Autoregression Estimation in Panel B***

	All MSAs			**Bubble MSAs**			**Non-Bubble MSAs**		
	2 & 4 lags	**2 & 4 lags**	**2 & 8 lags**	**2 & 4 lags**	**4 & 8 lags**	**4 & 8 lags**	**2 & 4 lags**	**4 & 8 lags**	**4 & 8 lags**
1990-1998	1.0315	1.0413	1.0095	1.0388	1.0609	1.0213	1.0012	1.0203	1.0191
1999-2006	1.1694***	1.1599***	1.0082***	1.2552***	1.2036**	1.0429***	1.0229	1.0348	1.0116
2007-2013	1.0064	1.0412	1.0346	1.006	1.0231	1.017	1.0768	1.1472**	1.0653
Period 1 & 2	1.3108***	1.4860***	1.4601***	1.3402***	1.6194***	1.5204***	1.5120***	1.5448***	1.5334***
Period 2 & 3	1.1150**	1.1428***	1.1151**	1.1354*	1.0996	1.0949	1.3572***	1.2618***	1.2071***
Period 1 & 3	1.4615***	1.6983***	1.6282***	1.1804**	1.4728***	1.3886***	2.0520***	1.9492***	1.8510***

Note : *, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

Table 4 Explanatory Power of Various Components in Model A1

	Mean	**Maximum**	**Minimum**	**Max - Min**
Non-bubble MSAs				
Long Run Partial R^2	0.1324	0.4437	0.0299	0.4138
Short Run Partial R^2	0.5145	0.8857	0.3006	0.5851
PMG Model	0.3034	0.7819	0.1301	0.6519
Residual Autregression	0.5035	0.7636	0.2059	0.5577
Total = PMG + Residual AR(4)	0.7345	0.9849	0.4137	0.5712
Bubble MSAs				
Long Run Partial R^2	0.2313	0.5533	0.0667	0.4866
Short Run Partial R^2	0.6472	0.9150	0.2506	0.6644
PMG Model	0.4198	0.7925	0.0902	0.7023
Residual Autregression	0.3635	0.7956	0.0293	0.7663
Total = PMG + Residual AR(4)	0.6790	0.9612	0.0464	0.9149

Table 5 PMG Estimations from Models B and C

Panel 5A PMG Estimation Results

	Model B				Model C	
	B1	B2	B3		C1	C2
Long run coefficients						
10-yr TIPs	0.1592***	0.1551***	0.1570***	Yield spread	0.0142***	0.0176***
10Y- RentG				10Y- RentG	0.1862***	0.2234***
10-year Treasury				10-year Treasury	0.1269***	0.1431***
Short run coefficients						
Error Correction	-0.0267***	-0.0266***	-0.0253***		-0.0203***	-0.0189***
$\Delta R/P_{t-1}$	0.0496	0.0432	0.2014***		0.0271	0.0322
$\Delta R/P_{t-2}$	0.0998***	0.1066***	0.0702***		0.0953***	0.0847***
$\Delta R/P_{t-3}$	0.1649***	0.1667***	0.1933***		0.1522***	0.1474***
$\Delta R/P_{t-4}$	0.1269***	0.1293***	0.1468***		0.1150***	0.1087***
ΔYield spread$_t$	0.0017***	0.0022***	0.0021***		0.0018***	0.0013**
ΔYield spread$_{t-1}$	0.0018	0.0017	0.0012**		0.0015	0.0016
ΔYield spread$_{t-2}$	0.0020*	0.0020*	0.0020***		0.0017	0.0017
ΔYield spread$_{t-3}$	0.0021***	0.0021***	0.0013**		0.0018***	0.0017***
ΔYield spread$_{t-4}$	-0.0001	0	0.0003		-0.0003	-0.0004
$\Delta 10Y_t$		-0.0018	0.3510***		-0.0039*	
$\Delta 10Y_{t-1}$		0.0082**	0.2062***		0.0065*	
$\Delta 10Y_{t-2}$		-0.0004	0.2571***		-0.0021*	
$\Delta 10Y_{t-3}$		-0.0026*	0.1091***		-0.0042***	
$\Delta 10Y_{t-4}$		-0.0003	0.035		-0.0016	
$\Delta 10Y_t$ - RentG$_t$	-0.0033		-0.3516***			-0.0058**
$\Delta 10Y_{t-1}$ - RentG$_{t-1}$	0.0089**		-0.1988***			0.0069*
$\Delta 10Y_{t-2}$ - RentG$_{t-2}$	-0.001		-0.2583***			-0.0029**
$\Delta 10Y_{t-3}$ - RentG$_{t-3}$	-0.0023		-0.1125***			-0.0041***
$\Delta 10Y_{t-4}$ - RentG$_{t-4}$	-0.0004		-0.0349			-0.0019*
Constant	0.0270***	0.0272***	0.0256***		-0.0229***	-0.0314***
Log Likelihood	12702	12707	13850		12726	12729
Hausman Test	-1.92	-27.69	0.36		**1.34**	**0.71**
p-value	**Invalid**	**Invalid**	0.5464		**0.7202**	**0.8711**

Note : *, ** and *** denote significance at the 10%, 5% and 1% levels respectively.
The variable 10-Y – RentG represent 10-year Treasury rates minus rent growth, yield spread is Merill Lynch 1-year high yield bond rates minus 1-year Treasury rates.
An insignificant value of Hausman Test indicates PMG is preferred to MG Estimation.

Appendix A Sum of Short-run Coefficients for Individual MSAs from PMG Estimation of Model A2

Panel A Non-bubble MSAs

Non-bubble MSAs	**Error Correction**	**sum of (Δ*R*/*P*)**	**sum of (ΔYield spread)**	**sum of (Δ10Y Treasury)**	**Constant**
Akron	-0.0303	0.2463	0.0077	0.0177	0.0352
Ann Arbor	-0.0353	0.4618	0.0000	-0.0107	0.0344
Atlanta	-0.0158	0.4285	0.0038	0.0009	0.0146
Atlantic City	-0.0201	0.6690	0.0090	0.0178	0.0243
Baltimore	-0.0252	0.8728	0.0048	-0.0004	0.0111
Boulder	-0.0252	-0.0868	0.0132	-0.0092	0.0161
Bremerton	-0.0179	0.5212	0.0078	-0.0158	0.0189
Chicago	-0.0281	0.7512	0.0059	0.0112	0.0366
Cincinnati	-0.0130	0.6262	0.0048	0.0150	0.0152
Cleveland	-0.0296	0.1864	0.0050	0.0188	0.0362
Dallas	-0.0107	0.2651	0.0041	0.0039	0.0095
Denver	-0.0182	0.4579	0.0054	0.0034	0.0145
Detroit	-0.0379	0.5713	0.0030	-0.0031	0.0411
Flint	-0.0117	-0.3830	0.0156	0.0012	0.0144
Fort Worth	-0.0120	0.1211	0.0040	0.0033	0.0117
Gary	-0.0351	0.3488	0.0102	0.0140	0.0515
Greeley	-0.0268	-0.0648	0.0036	-0.0017	0.0257
Houston	-0.0105	0.2116	0.0031	-0.0013	0.0069
Kansas City	-0.0294	0.0580	0.0031	-0.0014	0.0270
Lake County	-0.0164	0.4308	0.0042	0.0124	0.0237
Milwaukee	-0.0240	0.4724	0.0055	0.0127	0.0311
Minneapolis	-0.0255	0.3067	0.0050	-0.0095	0.0204
Philadelphia	-0.0208	0.6959	0.0031	0.0293	0.0258
Pittsburgh	-0.0338	-0.2266	-0.0008	0.0124	0.0361
Racine	-0.0170	0.5750	0.0047	-0.0053	0.0208
St.Louis	-0.0249	0.4735	0.0014	0.0029	0.0252
Wilmington	-0.0220	0.4458	0.0116	0.0110	0.0264
Average	-0.0229	0.3495	0.0055	0.0048	0.0242
Maximum	-0.0105	0.8728	0.0156	0.0293	0.0515
Minimum	-0.0379	-0.3830	-0.0008	-0.0158	0.0069

Panel B Bubble MSAs

Bubble MSAs	**Error Correction**	**sum of (Δ*R/P*)**	**sum of (ΔYield spread)**	**sum of (Δ10Y Treasury)**	**Constant**
Anchorage	-0.0345	-0.1519	-0.0053	-0.0113	0.0204
Boston	-0.0332	0.7903	0.0066	0.0059	0.0337
Fort Lauderdale	-0.0333	0.8314	0.0045	-0.0070	0.0272
Los Angeles	-0.0264	0.8309	0.0052	-0.0077	0.0239
Miami	-0.0356	0.3745	0.0090	-0.0103	0.0306
New York	-0.0251	0.8122	0.0054	0.0093	0.0287
Phoenix	-0.0380	1.0877	0.0004	-0.0022	0.0083
Portland	-0.0119	0.4946	0.0054	-0.0031	0.0121
Riverside	-0.0292	0.8154	0.0083	-0.0129	0.0275
Salem	-0.0247	0.7440	0.0061	0.0146	0.0334
San Diego	-0.0351	0.8033	0.0081	-0.0070	0.0379
San Francisco	-0.0239	0.7903	0.0087	-0.0052	0.0236
San Jose	-0.0234	0.7344	0.0122	-0.0133	0.0223
Seattle	-0.0200	0.6085	0.0107	0.0057	0.0200
Tacoma	-0.0294	0.6889	0.0035	-0.0046	0.0335
Tampa	-0.0324	0.9046	0.0040	-0.0001	0.0233
Honolulu	-0.1059	-1.2240	0.0055	0.1443	0.1587
Washington	-0.0267	0.8056	0.0050	-0.0037	0.0099
Average	-0.0327	0.5967	0.0057	0.0051	0.0319
Maximum	-0.0119	1.0877	0.0122	0.1443	0.1587
Minimum	-0.1059	-1.2240	-0.0053	-0.0133	0.0083

Appendix B Examples of Other PMG and MG Estimations

	Long Run is 10-yr Treasury – Rent Growth		Long Run is High Yield – Rent Growth	
	PMG	**MG**	**PMG**	**MG**
Long run coefficients				
Trend	0.0194***	-0.0095	-0.0059***	-0.0052***
10Y- RentG	0.2926***	-0.0523		
HY - RentG			-0.0366***	-0.0601**
Short run coefficients				
Error Correction	-0.0181***	-0.0289***	-0.0257***	-0.0352***
$\Delta R/P_{t-1}$	0.039	0.0269	0.0653*	0.0635*
$\Delta R/P_{t-2}$	0.0860***	0.0748***	0.1060***	0.1040***
$\Delta R/P_{t-3}$	0.1569***	0.1536***	0.1828***	0.1840***
$\Delta R/P_{t-4}$	0.1118***	0.1123***	0.1358***	0.1374***
ΔYield spread$_t$	0.0013**	0.0012**	0.0023***	0.0025***
ΔYield spread$_{t-1}$	0.0015	0.0014	0.0022*	0.0023
ΔYield spread$_{t-2}$	0.0015	0.0015	0.0024**	0.0027***
ΔYield spread$_{t-3}$	0.0015**	0.0016**	0.0025***	0.0027***
ΔYield spread$_{t-4}$	-0.0004	-0.0004	0.0006	0.0008
$\Delta 10Y_t$ - $RentG_t$	-0.0056**	-0.0056**	-0.0016	-0.0016
$\Delta 10Y_{t-1}$ - $RentG_{t-1}$	0.0072**	0.0070**	0.0104***	0.0102***
$\Delta 10Y_{t-2}$ - $RentG_{t-2}$	-0.0030**	-0.0029*	0.0002	0.0004
$\Delta 10Y_{t-3}$ - $RentG_{t-3}$	-0.0040***	-0.0038***	-0.0006	-0.0005
$\Delta 10Y_{t-4}$ - $RentG_{t-4}$	-0.0018	-0.0017	0.0012	0.0014
Constant	-0.0331***	-0.0074	0.0592***	0.0850***
Log Likelihood	12725	12808	12688	12740
Hausman statistic	**7.17**		**5.85**	
***P*-value**	**0.0277**		**0.0535**	