# House Price Efficiency: Expectations, Sales, Symmetry

**Arthur Grimes, Andrew Aitken, Suzi Kerr**

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## Author contact details
Arthur Grimes
Corresponding Author
Motu Economic and Public Policy Research
PO Box 24390
Wellington
New Zealand
Email: arthur.grimes@motu.org.nz

Andrew Aitken
Motu Economic and Public Policy Research
Email: andrew.aitken@motu.org.nz

Suzi Kerr
Motu Economic and Public Policy Research
Email: suzi.kerr@motu.org.nz

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## Motu Economic and Public Policy Research
PO Box 24390
Wellington
New Zealand

| | |
|---|---|
| Email | info@motu.org.nz |
| Telephone | +64-4-939-4250 |
| Website | www.motu.org.nz |

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# Abstract

An efficient housing market is of critical importance for individual welfare and for a well-functioning economy. We test the efficiency of this market by estimating the factors that determine both the long-run and the dynamic paths of regional house prices. Our tests use a new quarterly regional panel data set covering the 14 regions of New Zealand from 1981 to 2002. The tests indicate that regional housing markets converge to an equilibrium consistent with consumer optimising conditions, and hence with long-run efficiency. However, some conditions required for short-run (dynamic) efficiency are violated. We find that extrapolative price expectations, based on past regional phenomena, lead to overshooting of house prices in response to new region-specific information. We also find that price dynamics are influenced by past regional house sales activity and that the dynamic adjustment process is asymmetric depending on whether house prices are above or below their long-run equilibrium.



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G14, R21, R31

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House prices; housing appreciation; housing market; adjustment dynamics

# Contents



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# 1 Introduction

An efficient housing market is of critical importance for individual welfare and for a well-functioning economy (Di, 2001; Bajari and Kahn, 2003). Long-run efficiency requires that house prices converge to a relationship determined by consumers' optimising conditions. Short-run efficiency requires prices to adjust quickly to new information so that excess profit opportunities, after deducting trading costs, do not linger in the market.

We test for long-run and short-run efficiency of the housing market, as reflected in house prices. Our tests indicate that regional housing markets converge to an equilibrium consistent with long-run efficiency. However, some conditions required for short-run efficiency are violated. In particular, we find that extrapolative price expectations, based on past regional phenomena, lead to overshooting of house prices in response to new region-specific information. Further, price dynamics are influenced by past regional house sales activity. Notably, the dynamic adjustment process is asymmetric depending on whether house prices are above or below their long-run equilibrium. Thus the degree of short-run efficiency is affected by the nature of a shock's impact (upwards or downwards) on the equilibrium price.

Our tests use a new regional panel data set covering the 14 regions of New Zealand over 88 quarters [1981(1)–2002(4)]. This dataset includes the median sales price of owner-occupied dwellings, the ratio of the dwelling's sales price to its official valuation (used for property tax purposes), and the number of sales in each quarter. We use this data in conjunction with other relevant regional data to test the efficiency properties of the housing market. The sales data enables us to test for sales-driven "fad" (or other) effects that existing studies either cannot test or have to test using inadequate proxies. Our finding that sales activity has a material, but asymmetric, effect on house price dynamics adds to current understanding of the nature of property market dynamics.

We illustrate our key findings relating to inefficiency of house price dynamics through simulated house price paths in response to certain economic and financial shocks. Plausible economic developments, based on recent historical experience, indicate the potential for material house price overshooting.

House prices summarise a large amount of information regarding the desirability and cost of living in a particular location (MacDonald and Taylor, 1993; Case and Mayer, 1996; Sheppard, 1999; Cook, 2003) and influence migration patterns (Glaeser and Gyourko, 2001). Recent studies have cast doubt on the efficiency of this market. For instance, Capozza and Seguin (1996) find evidence of euphoria in metro house markets, while Case and Shiller (1988, 1989, 2003) find evidence of predictability in regional house prices that should not exist in an efficient market. Inefficiency in such a critical market may have important welfare consequences and may cause financial imbalances, especially following the bursting of a bubble, with material macroeconomic consequences.[1]

Case and Shiller (2003) define a housing bubble as "a situation in which excessive public expectations for future price increases cause prices to be temporarily elevated". They postulate that fundamental factors such as income growth and interest rates initiate a house price change, but expectations may then become self-reinforcing, setting a bubble in train. Their survey-based results indicate that house buyers in markets with recent high house price growth build in high expected capital gains for the following decade. This behaviour is consistent with an expectations-driven bubble. They also produce survey evidence that house market adjustment is likely to be asymmetric, with sellers preferring to withdraw from the market during a downturn rather than accepting the price consequences of selling in those market conditions. By contrast, in an upturn, house prices respond rapidly, albeit being curtailed longer term by new house construction.

[1] Concerns over such issues led *The Economist* magazine to devote its 31 May 2003 survey of property to this topic under the title "Close to bursting".

This paper builds on these insights, presenting a systematic examination of the sources of potential inefficiencies in regional housing markets. We estimate the factors which determine both the long-run and dynamic paths of regional house prices. We do so by adopting an optimising model of house price determination, which we use to estimate long-run house price determinants. Expectations are a key factor in this formulation. We then estimate clearly specified dynamic models, and test whether certain features discussed by Case and Shiller (1989, 2003) and others influence the dynamic behaviour of the housing market. These features may push prices temporarily away from equilibrium. Our analysis is at a regional level and incorporates both regional and national variables, as appropriate, within a panel cointegration framework. This approach complements the survey-based and single equation approaches of Case and Shiller (1989, 2003). By testing hypotheses within a rigorous theoretical and econometric framework, and with an alternative dataset incorporating sales variables not available to previous researchers, we are able to shed considerable light on the (efficient and inefficient) behaviour of the housing market.

The questions we address in order to assess long run and short run efficiency, include: Are prices in the long term driven by factors identified through a consumer optimization problem? Do expectations-driven fads have an impact and, if so, are these fads driven by regional or national developments? Does regional sales activity have an effect in fuelling regional fads, or is sales activity a stabilising factor driving prices towards equilibrium? Do the dynamic adjustment mechanisms incorporate non-linearities which may indicate information-based reasons for variations in price adjustment towards long run equilibrium? Do significant adjustment asymmetries exist and, if so, what do these asymmetries indicate about the nature of fads or related phenomena?

We address this set of questions using a new, consistently measured quarterly panel dataset for New Zealand. The country has a population of 4 million people spread over two main islands. With a combined land area similar to that of Japan or the United Kingdom, it is divided into 14 regions corresponding to Regional Councils.[2] Regional Council boundaries follow physical features—primarily major water catchments—so regions tend to be distinct economic entities; for instance, individual cities do not flow over council boundaries.[3]

The paper is arranged as follows. Section 2 outlines the theoretical basis for our long-run model and explicitly formulates the long-run and dynamic models to be tested. Data is described in Section 3, and estimation results are presented in Section 4. Interpretation of the dynamic results, focusing especially on the effects of capital gains expectations, sales activity and asymmetric adjustment processes on house price dynamics, is presented in Section 5. Section 6 presents a brief conclusion, with pointers to future work.

## 2 Theoretical model

Our theoretical approach to estimating regional house prices builds on the work of Pain and Westaway (1996), who formulated an optimising model to determine equilibrium house prices. We use this model as a basis for examining long-run efficiency of the housing market. We then incorporate error correction and dynamic adjustment aspects introduced by Capozza *et al* (2002) plus additional dynamic adjustment features to investigate aspects of short-run efficiency discussed above.

[2] There are 16 Regional Councils but, in keeping with our data sources, we amalgamate 3 small neighbouring councils into one (Nelson-Marlborough-Tasman).
[3] The physical distinctiveness of regions is reflected in the finding that New Zealand house prices do not share a single common trend, so national variables alone cannot explain regional house price developments (Grimes *et al*, 2003).

## 2.1 Long-run model

Pain and Westaway formulate the consumer problem as one where each household allocates its lifetime wealth over consumption of housing services ($c^h$; proxied by a constant, $\theta$, multiplied by the housing stock, h) and non-housing consumption (c) in each period of life and over its bequest. Use of a constant relative risk aversion utility function (with coefficient of relative risk aversion, $\gamma$) and aggregating over individuals results in the optimising equation explaining equilibrium real house prices (g) in (1):

$$\ln(g) = (1 - \gamma)\ln(\theta) - \gamma\ln(h) + \gamma\ln(c) - \ln(UC) \quad (1)$$

where:

- g is the ratio of quality-adjusted price of housing ($p^h$) to the price of non-housing consumption goods ($p^c$)
- h is the housing stock
- c is non-housing consumption
- UC is the real user cost of capital (discussed further below).

If c is unobservable (as it is with the regional data at our disposal), we can add an auxiliary hypothesis that c is determined as in (2):

$$\ln(c) = \alpha + \beta\ln(y) \quad (2)$$

where: y is an appropriate activity variable influencing regional non-housing consumption.[4]

[4] Note that c represents aggregate non-housing consumption, so y represents aggregate economic activity; no additional demographic scalar is required. If we were to make c also a function of UC in (2), it would not alter the nature of our final estimating equations, although interpretation of the UC coefficient would differ. (2) and subsequent equations also include error terms; these are assumed to have standard properties, but are suppressed for expositional purposes.

House prices are observed for bundles of housing and related services.[5,6] If (as in our case) house sales price data are not quality adjusted, we can add another auxiliary hypothesis linking the real unadjusted sales price ($p^u/p^c$) to the quality adjusted sales price as in (3):

$$\ln(p^u/p^c) = \ln(p^h/p^c) + \xi Z \tag{3}$$

where: Z is a vector of house-specific and locality-specific attributes and $\xi$ is an accompanying coefficient vector. Combining (1)–(3) yields:

$$\ln(p^u/p^c) = \delta - \gamma\ln(h) + \beta\gamma\ln(y) - \ln(UC) + \xi Z \tag{4}$$

where: $\delta = [(1 - \gamma)\ln(\theta) + \alpha\gamma]$.

## 2.2 Issues in implementing the long-run model

The stock of housing (h) in each area is determined jointly with house prices in the long run. The rate of change in the stock of houses will be influenced by factors such as costs of constructing new houses, the degree of vacant land available for housing and regulatory efficiency (Capozza *et al*, 2002; Case and Mayer, 1996; Glaeser and Gyourko, 2002). The housing stock changes only slowly over time and so can be considered a predetermined variable over short to medium time horizons. By contrast, the house price is an asset price and so is a "jump" variable, reflecting the influence of new information. We can therefore estimate (4) and identify the effect of changes in each of h, y and UC on long-run real house prices.[7]

[5] House-specific services have been proxied elsewhere through: number of bathrooms, lot size, fireplace, garage size, air-conditioning, basement, detached dwelling, patio and previous dwelling purchase price (Can, 1992; Dubin, 1992; Genesove and Mayer, 2001).

[6] Housing services also include amenity and location values, which have been proxied elsewhere through: neighbourhood quality index, land supply index, coastal situation, distance from city centre, school assessment scores, crime rate, per capita income and unemployment rate (Can, 1992; Capozza et al, 2002; Case and Mayer, 1996; Dubin, 1992; O'Donovan and Rae, 1997).

[7] Eventually h will adjust, so the long-run system-wide effect of a change in each explanatory variable on house prices has to incorporate the housing stock response. O'Donovan and Rae (1997) found that single equation estimates of aggregate New Zealand house prices gave very similar results to full system estimates, which included equations also for consumption and for housing investment.

The variables which enter Z may include some that are fixed over time but which vary cross-sectionally (e.g. latitude of the locality) and others which vary over time (e.g. changing house quality or changing amenities within a locality). The former set can be handled through the inclusion of fixed effects. The latter require region-specific proxies. These quality and amenity variables are likely to be slowly changing over time (e.g. the process of "gentrification" of an area, or changing attitudes towards a coastal location). If (as in our case) there is no comprehensive data to proxy for the latter elements of Z, we can capture the influence of these variables through inclusion of a quadratic time trend, with coefficients that are freely estimated for each region reflecting trends in region-specific attributes. Thus, if the quality of residential houses in one region is trending upwards relative to another region, the quadratic time trend can account for this and other (quadratically) trending factors.[8]

A key variable in (4) is UC, the real user cost of capital. In formulating this variable we note that, in New Zealand, owner-occupiers' mortgage interest payments are not tax deductible, nor are capital gains from housing taxed. If loan finance is the marginal source of finance for housing then there is no tax relief on the housing loan and no tax to pay on the housing services. Thus no tax rate should appear in the UC variable.[9]

The relevant real user cost facing a house purchaser is the real interest rate, r,[10] less the expected annual real capital gain on the house, $\dot{g}$. The real interest rate is identical across regions in any given quarter, but $\dot{g}$ may not be. As discussed by Case and Shiller (2003), the nature of capital gains expectation formation is of importance to any fad or overshooting effect.

[8] By restricting our attention solely to residential houses, we avoid change in housing quality caused by a shift from detached dwellings to apartments. In our dynamic equations, changes in composition of houses sold each quarter are also accounted for.

[9] If, however, other taxable investment opportunities constitute the marginal source of finance for funding house purchase then the tax rate should enter into UC, since the opportunity cost is taxed. Also, housing investors' interest payments are tax deductible. This leads to the problem that different investors face different tax rates and the relationship of these tax rates to each other has varied over time. Entering a single tax rate would not adequately capture the taxation effect for different individuals.

[10] We proxy r by the 90-day bank bill yield less the latest CPI inflation rate, each expressed in annual terms. Grimes (1994) finds that mortgage interest rates are set as a margin above the 90-day bank bill yield; the margin is incorporated into the equation constant.

The more backward looking expectations are, the more likely it is that prices can overshoot their fundamental values in response to a change in fundamental factors. We test four alternative proxies for $\dot{g}$.

First, building on O'Donovan and Rae's (1997) nationwide analysis of the New Zealand housing market, we set $\dot{g}$ equal to the past three years' annual real capital gain on houses at the national level.[11] The three-year horizon reflects expectations based on the medium-term national trend in real prices. We call the resulting user cost variable UC1. Given its construction, it is identical across regions in any given quarter.

Second, we set $\dot{g}$ equal to the past three years' annual real capital gain on houses at the regional level. The resulting user cost variable, UC2, differs across regions in any given quarter.

Third, we set $\dot{g}$ equal to the past year's annual real capital gain on houses at the regional level. The one-year horizon reflects expectations driven by shorter term factors, consistent with Case and Shiller's (1989) finding that one year's house price changes help predict the next year's price change. The resulting user cost variable, UC3, also differs across regions in any given quarter.

Fourth, we set $\dot{g}$ equal to zero, allowing the equation constant to proxy for a constant real capital gain expectation over time. We denote this variable as UC4. If there is no backward-looking element in expectations, this proxy should perform better than each of the other three proxies.

As in O'Donovan and Rae (1997), we note that each of the four alternative proxies approaches zero, and at times becomes negative during our sample. This makes it impossible to include ln(UC) as specified in (4). Instead, we include UC with a freely estimated coefficient.

[11] We take the average of the first and fourth years in constructing this variable to reduce noise. For instance, the 1985(1) observation is calculated as the annual rate of change between calendar 1981 and calendar 1984 average real house prices.

Another complication with UC in the New Zealand context is that interest rates were strictly controlled until 1984 by government regulation and are unlikely at that stage to have equilibrated the credit market. (The shadow cost of credit is likely to have been considerably above the actual interest rate, reflecting excess demand at the regulated rate.) To account for this factor, we enter UC starting from 1985(1), and supplement it with the inclusion of a dummy variable (UCD) taking the value of 1 prior to 1985(1) and 0 thereafter.

Taking each of the above factors into account, the long-run estimating equation for each region corresponding to our theoretical specification (using subsequent data terminology, with time subscript t) is:

$$P_{zt} = a_{0z} + a_{1z}Y_{zt} + a_{2z} H_{zt} + a_{3z}UCx_{zt} + a_{4z}UCD_t + a_{5z}TIME_t + a_{6z}TIME^2_t \quad (5)$$

where:

- $P_z$ is the log of the real median residential house sales price in region z [corresponding to $\ln(p^u/p^c)$]
- $Y_z$ is the log of real regional economic activity for region z [corresponding to $\ln(y)$]
- $H_z$ is the log of the residential house stock in region z [corresponding to $\ln(h)$]
- $UCx_z$ is the real user cost of capital (x = 1, 2, 3, 4 as described above)
- UCD is a dummy variable = 1 prior to 1985(1), 0 thereafter
- TIME is a linear time trend
- $TIME^2$ is the linear time trend squared.

$a_{0z}$–$a_{6z}$ are coefficients with $a_{1z}$, $a_{2z}$, $a_{3z}$ and $a_{4z}$ each constrained to be identical across regions, reflecting the coefficients in (4); $a_{0z}$, $a_{5z}$ and $a_{6z}$ are coefficients that reflect fixed and region-specific trend effects and so differ across regions.

For our estimates to be consistent with long-run market efficiency, (5) has to be a valid long-run equation across all regions. As discussed below, the stochastic variables included in (5) are non-stationary [I(1)]. A requirement for (5) to be a valid long-run equation, and hence for the equation to be consistent with long-run efficiency holding, is that the residual from (5) is stationary; if this were not the case, real house prices would not return to the consumer's optimising conditions following a shock. In particular, we require the residual to be stationary when $a_{1z}$, $a_{2z}$, $a_{3z}$ and $a_{4z}$ are each restricted to be identical across regions, since in each case the relevant coefficient reflects underlying parameter(s) that are hypothesised to be identical across regions.

## 2.3 Dynamic model

Market efficiency is most often tested through examination of price dynamics (Fama and French, 1988; Case and Shiller, 1989; Capozza & Seguin, 1996). As Capozza and Seguin demonstrate, however, short-run market efficiency cannot be tested solely through an examination of house price dynamics, since rentals also form part of the return to housing. Without information on rentals, we cannot interpret directly whether the dynamics of house prices are consistent with a no arbitrage condition.

Instead, we focus on the speed and nature of the adjustment of house prices to long-run equilibrium, defined in (5), noting—to anticipate the empirical results—that (5) passes the tests to be considered a valid long-run specification of house prices. In a perfectly flexible market with zero transactions costs and no information costs, house prices should adjust immediately to their long-run values consequent on a change in one or more of the explanatory variables in (5). Even if this were the case, stochastic errors could cause temporary deviations in the house price from long-run equilibrium, but in an efficient market these deviations should be fully unwound in the subsequent quarter.

These conditions can be specified within an error correction framework. Denote the equilibrium value of the log of the real unadjusted long-run house price [from (5)] as $P^*_{zt}$, and consider the error correction equation (6):

$$\Delta P_{zt} = b_{1z}\Delta P^*_{zt} + b_{2z}(P_{zt-1} - P^*_{zt-1}) + b_{3z}\Delta P_{zt-1} + b_{4z}X_{zt} \qquad (6)$$

Inclusion of $\Delta P_{zt-1}$, the lagged change in real house prices, allows for a partial adjustment mechanism.[12] $X_{zt}$ is a vector of other stationary variables that may potentially impact on the dynamics of real house prices in region z. Short-run market efficiency, as discussed above, requires $b_{1z} = 1$ for all z, $b_{2z} = -1$ for all z, $b_{3z} = 0$ for all z, and $b_{4z} = 0$ for all elements of X and all z.[13]

The specification in (6) is appropriate for an environment of costless information dissemination about fundamentals. Capozza *et al* (2002) examine the case where increased housing market activity improves information dissemination about the market price (and quality) of houses. In this imperfect information world the adjustment coefficients, $b_{1z}$ and $b_{2z}$ in (6), may themselves be a function of housing market activity. Our house sales data corresponds to a direct measure of housing market activity, unlike Capozza et al who did not have a direct proxy for such activity. Building on their approach, we model this imperfect information environment by allowing the serial correlation and reversion parameters in the dynamic equation for each region to be functions of a housing market activity variable specific to a region as in (7):

$$\Delta P_{zt} = b_{1z}\Delta P^*_{zt} + [b_{2z}+b_{2Az}(A_{zt} - \bar{A})](P_{zt-1} - P^*_{zt-1}) + [b_{3z} + b_{3Az}(A_{zt} - \bar{A})]\Delta P_{zt-1} + b_{4z}X_{zt} \qquad (7)$$

[12] The relative values of $b_{1z}$, $b_{2z}$ and $b_{3z}$ in each region determine the degree of lagged adjustment and/or overshooting behaviour relative to fundamentals. Capozza *et al* (2002) demonstrate that as the serial correlation coefficient, $b_{3z}$, increases, the amplitude and persistence of house price cycles tends to increase. As the absolute value of the reversion coefficient, $b_{2z}$, increases, the frequency and amplitude of the cycle tends to increase.

[13] A single exception to the latter requirement in our estimated equation is a freely estimated coefficient on a variable (COMP) which accounts for short-term measured price changes due to changing composition of house sales between quarters within each region. For instance, if a higher ratio of "good" houses sells in one quarter than in the previous quarter, we would expect to see measured sale prices rise (temporarily) from one quarter to the next.

In (7), $A_z$ is an independent variable influencing adjustment of house prices in each region and $\bar{A}$ represents the mean value of $A_z$. In this specification, for instance, a region that has a value of $A_{zt}$ greater than $\bar{A}$ will have faster reversion of prices to fundamentals than the mean speed of reversion if $b_{2Az}$ is negative.

In operationalising (7), there is an issue as to whether the mean value ($\bar{A}$) should be time invariant as postulated by Capozza et al. For variables that are trending over time, this specification would imply a gradual raising or lowering of the partial adjustment and reversion parameters. In some cases this may be economically sensible but in others it will not be. We consider that the most robust way of specifying the $A_z$ variable is to choose a form of the variable which does not trend significantly over the sample period, so that the sample mean is a stable baseline against which to measure deviation of actual movements from the norm. Our measure of $A_z$ is the ratio of house sales to the housing stock in each region.[14] If sales rise, we hypothesise that there will be improved information dissemination; hence we expect $b_{2Az} \leq 0$, which corresponds to faster reversion to long-run equilibrium (where $-1 < b_{2z} < 0$). If sales activity coupled with lagged price changes incorporates new and/or improved information, we would expect $b_{3Az} \geq 0$. Interpretation of these coefficients will indicate whether market efficiency is affected by information disseminated through house sales activity.

Sales activity should have no effect additional to that specified in (7) in an efficient market (i.e. if it is entered as a component of $X_{zt}$, its coefficient should be zero). However, it may be, as suggested by Case and Shiller (1989), that sales activity does affect price dynamics independently of the information reasons just outlined. To test if this is the case, we add current and lagged sales activity independently to the equation (affecting the constant term) as in (8). In (8) the current and lagged activity variables test whether sales activity has an independent effect on price adjustment over and above the interaction terms in (7).

[14] Capozza *et al* used population as an imperfect proxy for sales. However, house sales are more likely to capture dynamic effects than is a slow-moving variable such as population. We also have data (albeit for only half the full period) for building consents, reflecting forthcoming house construction activity. However, the partial coverage of this variable meant it was not statistically significant when included in our work and we do not discuss its role further here.

$$\Delta P_{zt} = b_{1z}\Delta P^*_{zt} + [b_{2z} + b_{2Az}(A_{zt} - \bar{A})](P_{zt-1} - P^*_{zt-1}) + [b_{3z} + b_{3Az}(A_{zt} - \bar{A})]\Delta P_{zt-1} + b_{4z}X_{zt} + \textstyle\sum_i b_{5iz}A_{zt-i} \quad (8)$$

The vector, $X_{zt}$, now excludes sales activity. It remains the case that short-run efficiency requires $b_{4z} = 0$ and $b_{5iz} = 0$ for all i.

Our final test of the dynamic structure of house prices is a test for asymmetric adjustment depending on whether the previous quarter's actual prices are above or below fundamentals. Glaeser and Gyourko (2001), Cook (2003) and Case and Shiller (2003) all indicate the potential importance of asymmetric adjustment for the dynamics of the housing market. If regional housing demand expands, say in response to an increase in regional economic activity, prices are expected to rise from (5) but housing supply (h) will also expand gradually over time. By contrast, if regional housing demand falls, housing supply is unlikely to contract materially other than through depreciation (Glaeser and Gyourko, 2001). The effects of these asymmetric factors on expectations may be reflected in asymmetric adjustment to equilibrium in the two situations, with prices reacting more strongly to a fall in equilibrium prices than to a rise. Another cause of asymmetry (working in the opposite direction) may be reluctance by previous buyers to experience a realised capital loss in situations where prices have fallen. Sales may reduce in such circumstances without a significant observed price fall (Genesove and Mayer, 2001).

We test whether coefficients in (8) are identical when the sample is split into two categories: $P_{zt-1} > P^*_{zt-1}$ and $P_{zt-1} < P^*_{zt-1}$. If identical, then adjustment is symmetric; otherwise asymmetric adjustment is indicated. Asymmetric adjustment may indicate inefficiencies under some market conditions, as in the "reluctant-seller" (Genesove and Mayer) case cited above. Interpretation of the coefficients in the split sample will yield insights into what is driving any asymmetry.

# 3 Data

We use Quotable Value New Zealand (QVNZ) data for median residential house sale prices in each region. QVNZ is a state-owned entity that collects data on all house sales and which also values properties for local authority property tax purposes. We have measures, from this source, of the number of house sales in each region, the QVNZ valuation of houses that are sold, and the median sales price. Each of these variables is used in the estimation in Section 4. In order to compare "like with like" as much as possible, we restrict our attention to the residential house market, which excludes all multi-unit residential sales and all non-residential transactions. All data is available quarterly from 1981(1) – 2002(4). These data, together with data for the regional housing stock, are described in detail in Grimes et al (2003). That paper also presents tests for bivariate cointegration between regional house price levels and bivariate contemporaneous correlation between regional house price changes. These tests indicate that while house prices are cointegrated for some regional pairs they are more frequently not cointegrated. A little over half the contemporaneous correlations are significant at the 5% level. Together, these results indicate some similarity in house price developments nationally, but also reveal material elements of regional diversity.[15]

Regions are denoted RC01-RC15 (there is no region 10); RC01-RC09[16] are in the North Island, RC11-RC15[17] are in the South Island. Table 1 lists key characteristics of the data for each region. Column 1 presents the median nominal sales price for the 2002 calendar year, demonstrating that the median price in Auckland (RC02), New Zealand's largest city, was 4.5 times that in the (rural) West Coast of the South Island (RC12). Column 2 presents the change in real sales price between 1981 and 2002 (i.e. after deflating the median sales price by the consumers price index, CPI).

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[15] Grimes *et al* also conduct Granger causality tests on regional house prices for all regional pairings, in both directions. Three-quarters of the 182 tests are not significant at the 10% level, implying that spatial autocorrelation is not material at the regional council level.

[16] Northland, Auckland, Waikato, Bay of Plenty, Gisborne, Hawke's Bay, Taranaki, Manawatu-Wanganui and Wellington respectively.

[17] Nelson-Marlborough-Tasman, West Coast, Canterbury, Otago and Southland respectively.

Real prices in Southland (RC15) fell by 27% over this period, while those in Gisborne (RC05) were virtually unchanged; both regions are predominantly rural with no city having a population in excess of 47,000. By contrast, real prices in Auckland more than doubled and those in Wellington (RC09), the capital city, almost doubled. The average number of quarterly sales throughout the sample for each region is shown in column 3; column 4 lists the population of each region at the 2001 census; column 5 presents population density at the 2001 census.

In order to proxy real regional economic activity ($Y_{zt}$), we use the logarithm of the National Bank of New Zealand Regional Economic Activity indices (National Bank of New Zealand, 2003). Column 6 of Table 1 indicates the percentage change between 1981 and 2002 in this variable for each region. As with previous columns, considerable divergence in regional performance is indicated. Growth in the fastest growing region (Northland, RC01) was over three times that in Gisborne (RC05). A comparison of columns 2 and 6 indicates that fast-growing regions tended to have faster-growing real house prices; the (cross-sectional) correlation between the two columns is 0.72.

In the dynamic equations we need to account for changes in the composition of houses sold within a region in a particular quarter. To do so, we use the QVNZ valuation (as opposed to sales price) data for the houses sold in each quarter in each region. We form a composition variable, $COMP_z$, which takes the ratio of the median valuation of houses sold in a region relative to a Hodrick-Prescott filtered series for that region's median house valuation, the latter series representing the trend valuation of houses in the region. If the ratio in a quarter is greater (less) than one, the median house sold in that quarter is better (lower) quality than the average house in that region. Hence this variable should enter the dynamic equation with a positive sign.[18] The sales variable, $S_z$, which we enter into the dynamic equation (representing the housing market activity variable, $A_z$) is the ratio of house sales to the housing stock in each region. House sales data are obtained from QVNZ.

[18] This variable is stationary and so does not appear in the long-run equation.

Each of Pz, Yz, Hz and UCxz are tested for non-stationarity using the panel unit root tests of Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2002). We also test COMPz and Sz, which are each included in the dynamic specification. Results are presented in Table 2. Where the results of these tests are unambiguous (i.e. consistent from the 1% through to the 10% significance level) the implied order of integration is indicated in Table 2; ambiguous results are shown as I(0)/I(1).

Each of the variables, other than $COMP_z$, is either unambiguously non-stationary or else the order of integration cannot be determined with (near) certainty. Where a result is ambiguous, we prefer to treat the series as non-stationary, unless theory suggests that stationarity is more appropriate.[19] Each of the stochastic variables in (5) is therefore treated as being I(1).[20] $COMP_z$ is clearly I(0); $S_z$ is also treated as I(0) since the sales to house stock ratio must be bounded above and below, indicating stationarity.

# 4 Results

## 4.1 Long-run results

We estimate (5) using both OLS and SUR, presenting results for each estimation method.[21] Initially we estimate the equation with no restrictions. To provide a basis for comparison with our subsequent estimates, the Adjusted $R^2$ and standard error (s.e.) for the system, using UC2, are 0.9988 and 0.0445 respectively.[22] The s.e. indicates an average error of 4.5% across the sample, which can be compared with an s.e. of 7.9% when the system is estimated with just the fixed effects and quadratic time trend terms included.

We test the system of unrestricted equations for cointegration using the group mean panel test of Pedroni (1999).[23] The test statistic is –11.88 against a

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[19] See Banerjee *et al* (1993).
[20] ADF tests on UC1 and UC4, both national variables, indicate that they are I(1) with drift.
[21] Estimation is done in Stata and in Eviews. The Stata OLS results correspond to the Eviews WLS results. The SUR results are identical in each.
[22] As shown in Table 3, UC2 provides the greatest explanatory power of all the UC variables.
[23] This parametric ADF-based test is analogous to the Im *et al* unit root statistic applied to the estimated residuals of a cointegrating regression. The test allows for heterogeneity in both the

critical value of –1.46 indicating that the system of equations is cointegrated. Thus the unrestricted estimates are consistent with a valid long-run specification.

When we restrict each of the coefficients $a_{1z}$, $a_{2z}$, $a_{3z}$ and $a_{4z}$ to be identical across all regions, the Adjusted $R^2$ and s.e. for the system are 0.9980 and 0.0485 respectively, little changed from the unrestricted estimates. Application of the Pedroni panel cointegration test yields a test statistic of –8.33 against a critical value of –6.28, which indicates that the restricted system is also cointegrated and so represents a valid long-run specification for regional house prices. In terms of efficiency, this finding implies that the housing market is consistent with long-run efficiency, whereby house prices converge to values consistent with consumer optimisation.

Table 3 presents the restricted OLS and SUR system estimates using each of UC1–UC4. The results are used to examine the nature of the expectations process for house prices. In each case a constant TIME and $TIME^2$ are included unrestricted for each region in addition to the four variables listed, but are not reported in the table for clarity. The Adjusted $R^2$ and s.e. for the system are included for each estimation for comparison purposes.

The results are similar across the two estimation techniques, and show consistency in sign and broad magnitude of coefficients across the different UC specifications. However, the explanatory power of the equations differs substantially across the different specifications of UC. By far the strongest explanatory power comes with UC2, embodying region-specific real capital gains expectations based on medium-term (past three-year) developments.

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long-run cointegrating vectors as well as heterogeneity in the dynamics associated with short-run deviations from these cointegrating vectors.

The s.e. for each of the national specifications and for the one-year region-specific specification of UC are all similar, and each is over 10% higher than for the corresponding UC2 estimates. The coefficient estimates on UC are considerably higher in absolute value for UC2 than for the other UC specifications and the significance of those estimates is very much higher than for the other three UC specifications. In the dynamic equations that follow, the greatest explanatory power is also obtained using UC2 ahead of any of the other UC specifications. We therefore restrict our attention to this definition of UC in the remainder of the paper. The implications of the three-year region-specific UC specification for house price dynamics are analysed in Section 5.

Restricting attention to the UC2 specification, the elasticity of real house prices with respect to regional economic activity is approximately unity (0.92 using SUR; 1.17 using OLS). As the housing stock expands, *ceteris paribus*, the real price of housing falls with an elasticity of approximately two-thirds. Each of these estimates appears intuitively reasonable.[24] In most regions, the coefficient on TIME is positive while the coefficient on $TIME^2$ is negative. At the end of the sample, the combined coefficient effect of TIME and $TIME^2$ on $P_z$ was positive in ten of the fourteen regions.

A one percentage point increase in the real user cost of capital is estimated to reduce the long-run real house price by between three-quarters of one per cent and one per cent. Given that the real user cost of capital does not change markedly over long periods, the long-run upward trend in real house prices in most regions cannot be attributed to financial factors; rather, the upward trend in real house prices is attributable mainly to increases in economic activity and to the effects of tastes and other factors proxied by TIME and $TIME^2$.

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[24] Compared with the underlying structural parameters in (4), the SUR estimates indicate a CRRA ($\gamma$) of 0.65 and a consumption elasticity with respect to economic activity ($\beta$) of 1.43. As discussed in Grimes *et al* (2003), our measure of the housing stock may involve some inaccuracy which could lead to the absolute value of the estimate for $\gamma$ being understated, and hence to the implied estimate for $\beta$ being overstated. (Any trend error in the housing stock estimate will be compensated for by inclusion of the quadratic time trend for each region.)

The positive coefficient on UCD indicates that real house prices were higher, *ceteris paribus*, before financial deregulation than afterwards. The regulated period was one in which real interest rates were frequently negative (boosting house prices) while the supply of credit was restricted (reducing house prices). The positive coefficient on UCD indicates that the former effect outweighed the latter over the 1981–1984 period.

## 4.2 Dynamic results: Symmetric

To test short-run efficiency, we start with the symmetric dynamic framework in (8) with the interaction terms ($b_{2Az}$ and $b_{3Az}$) constrained to zero. The impact of the interaction terms is tested subsequently.

Recall that short-run efficiency requires that $b_{1z} = 1$, $b_{2z} = -1$, and requires all other coefficients, other than the coefficient on $COMP_{zt}$ (accounting for the effect of compositional changes in house sales on the measured median price), to equal zero. We split $\Delta P^*_{zt}$ into its individual components ($\Delta Y_{zt}$, $\Delta H_{zt}$, $\Delta UC2_{zt}$) to test the short-run adjustment speed for each of its constituent parts. In this case, the requirement that $b_{1z} = 1$ corresponds to a requirement that the coefficients on each of these components equal their long-run counterparts from Table 3.

In the $X_{zt}$ vector, we include $COMP_{zt}$, as described above. We also include the log change in consumer prices, $\Delta PC_t$. Inclusion of this variable allows us to test whether aggregate consumer price changes are fully and immediately incorporated into regional house price changes. If this is the case, the coefficient on $\Delta PC_t$ will be zero (i.e. $b_{4z} = 0$). A coefficient between –1 and 0 indicates some measure of partial adjustment of house prices to consumer price changes, in which case changes to consumer price inflation will impact temporarily on real house prices.

From (8), we include current and lagged sales activity, $S_{zt-i}$, as our measure of housing market activity. This variable tests for "bandwagon" or other effects of sales activity on prices that arise separately from the information dissemination role of sales posited by Capozza et al. In an efficient market, the coefficients on these (current and lagged) sales variables should equal zero.

In estimating the specification based on (8), the lagged residual was highly significant, but no separate partial adjustment process for nominal house prices was found significant (i.e. $b_{3z} = 0$). Thus lagged house price changes ($\Delta P_{zt-1}$) are omitted from the reported results in Table 4. Cross-equation restrictions are again imposed on the system. The resulting OLS equation is presented as column (1) in Table 4, where $RES_{t-1}$ is the lagged residual, using UC2 from the OLS equation presented in Table 3. Inspection of the estimates in column (1) reveals a number of features that relate to our tests of short-run efficiency.

First, the coefficient on the lagged residual ($b_{2z}$) is significantly negative; its high t-value (18.73) confirms the cointegration findings from Table 3, indicating also that the preferred equation from Table 3 is a valid long-run equation explaining $P_{zt}$. However, the coefficient is significantly different from –1, with the 95% confidence interval being (–0.48, –0.39). This estimate indicates that any deviation in house prices from equilibrium in one period is not fully unwound in the subsequent quarter. Further, the coefficients on $\Delta Y_{zt}$ and $UC2_{zt}$ are well below, and significantly different from, their long-run counterparts; the coefficient estimate on $\Delta PC_t$ (which is significantly different from zero) indicates that around half of consumer price inflation is reflected in house prices contemporaneously. The only coefficient that is not significantly different to its long-run counterpart is that on $\Delta H_{zt}$.

Together, these results indicate that we can reject short-run efficiency in the sense that prices do not adjust to existing disequilibria or to short-run shocks within one quarter. Nevertheless, the adjustment parameters are highly significant and indicate that approximately half the adjustment to most shocks occurs within a one quarter timeframe.

Thus, heuristically, the degree of short-run inefficiency appears small, especially for a market that does not have freely traded liquid securities.

When including current and lagged sales activity, $S_{zt-i}$, we tested for lags individually and also tested whether a group of current and/or lagged variables was significant. When $S_{zt-2}$ is included, no other lag of sales is significant, and $S_{zt-2}$ always outperformed any other lag of that variable. Thus we include $S_{zt-2}$ as our measure of housing market activity. Its coefficient is particularly interesting in terms of testing for short-run efficiency. Commonly, it is observed that there is a correlation between sales activity and house prices, but our results suggest that after controlling for other influences on house prices, current sales activity has no additional explanatory power. Instead, sales activity influences house prices with a two-quarter lag. The significant positive value for this coefficient is contrary to short-run market efficiency. A reasonable interpretation is that prospective buyers see housing activity lift, decide to embark on house purchase/sale; there is then a four- to six-month lag between their observing the housing market activity and their actual market involvement. This behaviour significantly affects the dynamics of house prices.

The final variable in the equation, $COMP_{zt}$, is highly significant, indicating that the composition of houses sold within a quarter affects the median price observed within a region that quarter. Thus adjusting for this effect is important given the sales data that we have. The significant positive coefficient on $COMP_{zt}$ is not an indicator of market inefficiency.

We have run a number of checks on the robustness of the parameters reported in column (1). Column (2) of Table 4 estimates the same equation using SUR, using the SUR long-run specification (with UC2) from Table 3. There is little change to any of the coefficients or to the explanatory power of the equation.

Tests of the residuals from the OLS equation in Table 4 indicate the presence of autocorrelation and heteroskedasticity. In column (3), we present estimates with panel-corrected standard errors using the Prais-Winsten (PW) method;[25] in column (4) we present generalised least squares (GLS) estimates.[26] In each case, there is little difference in coefficient estimates.

## 4.3 Dynamic results: Asymmetric

The potential for asymmetric adjustment was discussed in relation to (8). Estimates of asymmetric adjustment may be particularly useful in exposing the circumstances contributing to the short-run inefficiency found above; for instance, adjustment may be less consistent with market efficiency when the market is either above or below equilibrium.

In order to investigate the potential for asymmetric adjustment, we re-estimate column (1) in Table 4, with a dummy term (equal to one when the lagged residual is positive and zero otherwise) that is interacted with each of the variables in the equation. This allows us to estimate separate coefficients on each variable depending on whether the lagged residuals are positive (prices above equilibrium) or negative. The results of estimating this asymmetric adjustment process, using OLS, are shown as column (5) in Table 4.[27]

Several features stand out in these results. First, adjustment to lagged disequilibrium is estimated to be very much faster than in the symmetric case. The estimated coefficient is almost identical across negative and positive residuals; we cannot reject identical coefficients at the 5% significance level. In each case, however, the coefficient is still significantly different from –1.

[25] In the Prais-Winsten regression the disturbances are assumed to be panel-level heteroskedastic in the presence of first-order autocorrelation where the coefficient of the AR(1) process is specific to each panel.

[26] The GLS estimates allow for a heteroskedastic error structure with AR(1) autocorrelation specific to each panel. The adjusted $R^2$ and s.e. are not available for this option.

[27] SUR estimates provide almost identical results and so are not presented here.

Second, house prices adjust symmetrically to contemporaneous economic activity developments (statistically, we cannot reject symmetry). The estimated speed of adjustment to economic activity developments is materially stronger than in the symmetric case, and now represents approximately 80% of the estimated long-run response. The upper end of the 95% confidence interval in each case (1.11 and 1.07 with negative and positive residuals respectively) falls just short of the estimated long-run parameter.

Third, additions to the housing stock have a highly asymmetric effect on prices. In a depressed housing market, i.e. when house prices are below equilibrium (negative residual), additions to the housing stock have no effect on the price; the softening effect of new housing stock on house prices is felt only at times when the market is buoyant (prices above equilibrium). It is likely that the house stock expands principally in buoyant rather than depressed times. Thus in depressed times there is little explanatory power of housing stock changes on price (hence the parameter estimate which is not significantly different from zero). The price response to house additions in buoyant times is not significantly different from the estimated long-run response.

Fourth, the response of house prices to user cost changes is estimated to be virtually identical in buoyant and depressed conditions. The estimate is considerably higher than in the equation where symmetry is imposed on all coefficients and now represents almost 75% of the estimated long-run response. As in the economic activity case, however, the upper end of the 95% confidence interval falls just short of the estimated long-run parameter.

Fifth, consumer price inflation is estimated to be fully incorporated into house prices contemporaneously under buoyant market conditions; under depressed market conditions only three-quarters of consumer price inflation is contemporaneously embodied in prices. Symmetry in this respect can be rejected at the 5% level.

Sixth, the effect of lagged sales activity is also highly asymmetric. When the housing market is depressed, a rise in sales activity boosts prices, whereas in a buoyant market, a rise in sales activity has no statistically significant effect on house prices. This result is important for interpreting the role of sales in creating, or adjusting to, fads. A reasonable interpretation is that buyers and sellers are aware of market conditions in a buoyant market situation. In a depressed market people may hold off house purchase (for whatever reasons) but once people see the housing market starting to move (by observing increased sales) they enter the market intending to purchase a house prior to prices reverting to equilibrium. In this case, the sales variable can be considered an equilibrating factor by speeding the return of prices towards fundamental values within a depressed market situation. Nevertheless, even though it may be an equilibrating factor in this respect, its statistical significance is still indicative of the presence of short-run inefficiency at times in the housing market.

Finally, the explanatory power of the equation is considerably higher than in the symmetric case. The equation standard error falls by 27% and the Adjusted $R^2$ more than doubles with the asymmetric estimates relative to the symmetric case. This material change in explanatory power indicates that significant asymmetries in the adjustment process exist. When the asymmetric equation is estimated as two separate equations (split according to the sign of the residuals), the explanatory power is almost identical across depressed and buoyant market conditions (with an s.e. of 0.0267 and 0.0265 respectively). Thus house price changes are equally explicable in depressed and buoyant conditions, but material differences are found in the role of some variables, especially housing stock changes, consumer price inflation and sales activity. A comparison of column (1) and column (5) in Table 4 indicates that the linear model needs to incorporate asymmetric adjustment. (5) is therefore our preferred short-run, linear equation.

## 4.4 Dynamic results: Non-linear

The short-run estimates have hitherto constrained the dynamics to be linear (albeit asymmetric). They have so far ignored the potential for non-linear adjustment embodied in coefficients $b_{2Az}$ and $b_{3Az}$ in (8). The significance of the sales variable in the previous dynamic specifications suggests it is particularly important to investigate potential non-linearities associated with improved information dissemination arising from increased sales activity. We do so initially based on the symmetric specification reported as column (1) in Table 4, and then with asymmetries. In each case, we ignore the role of $b_{3Az}$ since we detect no significant role for lagged price changes in the adjustment process.

In implementing (8) we must decide whether to set $\bar{A}$ as the national mean of the housing market activity variable or as the regional mean, the latter varying across regions. It is possible, for instance, that information dissemination may be region-specific, in which case the latter variable will be more relevant, but if information dissemination is nationwide the national mean is relevant.

We estimate both regional and national versions of (8). In each case, we use the sales variable, $S_{zt}$, for our measure of housing market activity. The results from using the national and regional means are almost identical and there is little to choose statistically between specifications. Coefficients on each of the variables within the equation (including the lagged sales variable) are hardly altered by the alternative measures. The interaction term ($b_{2Az}$) is statistically significant in each specification ("t-values" for the national specification are 3.36 and 3.58 for OLS and SUR respectively; and for the regional specification they are 2.92 and 2.68 respectively).

The $b_{2Az}$ coefficient in each case is positive, implying that a higher sales ratio leads to slower adjustment to disequilibrium. The effect, however, is small. A 10% increase in $S_{zt}$ is estimated to alter the adjustment term on the residual from –0.414 to –0.379, implying little material shift in adjustment to disequilibrium. (This estimate is based on the OLS national specification; other specifications are similar.)

The direction of this result contrasts with the information-based reason put forward by Capozza et al for including the interaction term, since higher sales activity should improve information dissemination, therefore leading to faster adjustment to disequilibrium. An alternative explanation, consistent with the results here, is that high current sales activity has some fad element associated with it. For instance, in a buoyant market, high sales activity may delay adjustment back to fundamentals when price is above equilibrium.

We shed light on this potential explanation by re-estimating (8), using the national mean, with asymmetric adjustment depending on whether lagged residuals are positive or negative. We find that the interaction term is not statistically significant when the market is depressed but is just significant (at the 10% level in each of the OLS and SUR approaches) in a buoyant market. This result suggests that high sales activity has a slight delaying effect on adjustment to fundamentals in a buoyant market. Again, however, the effect is not material in an economic sense (a 10% increase in $S_{zt}$ in a buoyant market reduces the adjustment term on the residual from –0.697 to –0.673). Given the lack of materiality of the non-linear adjustment process—both symmetric and asymmetric—our interpretation of results henceforth concentrates on the linear asymmetric results, i.e. column (5) of Table 4.

# 5 Interpretation

To interpret our results and to apply them to recent housing market experience internationally, we examine the potential for overshooting or other fad-like (bubble) phenomena to arise consistent with our estimates. Taking the long-run (Table 3) and dynamic asymmetric OLS estimates (column 5 in Table 4) as our starting points, we trace out the dynamic effect of a realistic change to real economic activity on the real median house price. Initially we hold all other factors constant (i.e. we do not consider any flow-on effects to sales activity, etc) other than expectations which follow the extrapolative form estimated (within UC2) in the long-run equation. Subsequently, we examine the effect of interactions with house sales.

The economic activity change that we consider is based on actual aggregate US GDP experience since 1985.[28] In the decade to 1995q2, real GDP grew at an average rate of 0.72% per quarter (p.q.). Over the following five years (to 2000q2) the average growth rate was 1.05% p.q.; in the following two years (to 2002q2) average growth fell to 0.25% p.a. We take, as our baseline, a GDP (economic activity) growth rate of 0.72% p.q. and calculate the real house price corresponding to this track, holding other variables constant. We then compare the house price arising from a "shocked" GDP track and express this latter track as a percentage of baseline. The shocked track that we compare is one that historically grows at 0.72% p.q.; then (from quarter 1) experiences growth of 1.05% p.q. for 20 quarters, then 0.25% p.q. growth for the following 8 quarters, thereafter returning to 0.72% quarterly growth.[29]

Figure 1 graphs the resulting real house price path expressed as a percentage of baseline. The GDP track results in the long-run level of GDP in the shocked case settling 2.9% higher than baseline. This has the effect of raising long-run real house prices by 3.4% relative to baseline. In the interim, however, house prices rise to a peak at 8.2% above baseline (after 20 quarters) before dropping to 3.1% above baseline (after 35 quarters), thence returning to the long-run value.

In part, this behaviour is driven by the faster then slower path for GDP growth. But it is also affected by the expectations adjustment mechanism that feeds into the user cost variable. This mechanism can be seen from Figure 2. This figure graphs the real house price path consequent on a permanent 1% innovation to economic activity. The long-run house price effect of the activity increase is 1.17%; the contemporaneous effect is 0.89%. The initial house price increase feeds into real capital gains expectations via the UC variable so that UC falls consequent to the house price rise.

[28] US GDP data underlying these calculations is sourced from Bureau of Economic Analysis, US Department of Commerce.

[29] While this period involves quite a stark cycle for the US, we note that individual regional economies, as discussed by Case and Shiller (2003), have undergone considerably larger economic activity cycles with the potential for considerably magnified housing cycles relative to the aggregate experience.

The fall in UC is magnified as the house price rise continues, and this fall contributes to a further rise in the house price. The combined effect of this process and the direct effect of regional activity on house prices is to cause an overshooting of house prices above the long-run equilibrium. At some point, the positive residual created through house prices being above equilibrium exerts downward price pressure, equilibrating the market. A damped cycle results, as depicted in Figure 2. The peak of the cycle occurs in the 13$^{th}$ quarter, with a price rise (above baseline) of 1.24%. The trough occurs in the 26$^{th}$ quarter, just below the long-run value.

Endogenous sales activity is a factor which may make for additional complexity in the dynamics. Sales activity may respond to the price dynamics, i.e. to the change in prices and/or to the disequilibrium in prices. If that is the case, the presence of a significant sales effect in the dynamic price equation will affect the price path in response to a shock. To illustrate the effect of sales, we estimate a simple equation for $S_{zt}$ as a function of current and lagged changes in real house prices ($\Delta P_{zt-i}$) and the lagged residual from the long-run house price equation ($RES_{zt-1}$). Coefficients on these terms are restricted across regions, but constant and quadratic time trend terms are included and vary across regions to allow for region-specific fixed and trend effects. Estimates are presented in Table 5.

The estimates in Table 5 indicate that sales activity increases with current and lagged (real) price rises. The effect is strongest as a result of one quarter lagged price changes, with a decreasing effect thereafter, up to five quarters. This effect may contribute to a disequilibrating dynamic given that lagged sales in turn positively affect price changes. However, sales activity is also estimated to decrease as prices rise beyond their long-run value via the negative coefficient on the residual term. This has an equilibrating effect on the market given the lagged sales effect on prices.

Combining the dynamic equation for sales with the long-run and dynamic (asymmetric) price estimates, we calculate the price effect of an activity rise allowing for the interaction of direct price effects, the UC effect and the sales effect.

In the case of a 1% permanent increase in economic activity, there is a slightly greater degree of price overshooting than observed previously (with exogenous sales activity). The speed of the overshooting occurs considerably faster, with the peak being reached in the fourth quarter. (Apart from the speed of the overshooting, the other properties of the cycle remain similar to those shown in Figure 2 and so are not reproduced here.)

In response to a permanent 1% reduction in activity, a much slower cycle is exhibited with endogenous sales, with the trough in prices occurring in the ninth quarter. The reason for this asymmetry is the asymmetric adjustment of prices to sales. With a positive shock, the contemporaneous price rise is initially less than the long-run rise, resulting in a negative residual. The negative residual in turn causes a sales rise in excess of that driven by the initial price change, and the compound rise in sales further boosts prices (with a two-quarter lag) contributing to the speed of the overshooting. By contrast, a negative shock results in a positive residual (as prices initially fall less than the long-run fall) but in this case there is no material price response to sales because of the asymmetric effect of sales on price dynamics. Thus the effect of the shock takes longer to feed through to prices and overshooting occurs more gradually. This asymmetry mirrors the findings of Genesove and Mayer's (2001) "reluctant-seller" case whereby a negative localised shock leads to slower house price adjustment to equilibrium than is the case with a positive shock.

Overall, our estimates indicate that realistic changes in the path of economic activity can have a material effect on house prices, causing prices to overshoot their long-run equilibrium. Both extrapolative expectations effects and sales dynamics are shown to impact on the cycles that emerge from changes to underlying economic factors, while asymmetries in adjustment may be material.

# 6 Conclusions

Consistent with our theoretical model based on consumer optimisation conditions, regional real house prices converge to a long-run equilibrium determined by regional economic activity, the regional housing stock and the user cost of capital. Trend variables, reflecting trends in housing services and amenity values, have a significant impact in eleven of the fourteen regions.

In contrast with the long-run results, the strict conditions for short-run efficiency are not met. House prices respond to contemporaneous shocks in the directions expected and respond significantly to lagged disequilibrium in prices. However, adjustment is not fully completed within one quarter either to contemporaneous shocks or to lagged disequilibrium. Adjustment is symmetric in response to lagged disequilibrium and to contemporaneous shocks to economic activity and to the user cost of capital, but is asymmetric in response to consumer price changes. Further, house prices respond positively to past house sales activity, but only in depressed market conditions (when prices are below equilibrium).

The presence of a significant positive sales effect would normally be considered to contribute to price overshooting following a shock to fundamentals (for example, to economic activity). This is especially the case when sales activity is itself a positive function of past house price changes, as indeed we find. In the case of our estimates, however, the sales effect is an equilibrating influence, only driving prices upwards (towards equilibrium) when prices are below their equilibrium levels. When prices exceed equilibrium, the sales effect is absent.

Our estimates indicate that extrapolative house price expectations, based on medium-term regional price trends, provide the best explanation (amongst the expectations proxies we examined) of house price developments when incorporated into the user cost of capital measure. Our simulations of house price responses to economic activity shocks demonstrate that this expectations process induces some mild overshooting of house prices following an economic activity shock. These results are consistent with the survey-based findings of Case and Shiller (2003).

The length of the cycle depends on whether the activity shock is positive or negative (the latter having the longer cycle), reflecting the asymmetric influence of house sales activity on real house prices. The latter effect is consistent with sellers being reluctant to sell their houses when equilibrium prices have fallen following a negative economic shock.

Although our results reject short-run efficiency, heuristically the degree of housing market inefficiency appears small. Approximately three-quarters of lagged disequilibrium disappears within three months. A similar fraction of the long-run price effects from economic activity and user cost of capital shocks are reflected contemporaneously in house prices. Three-quarters of consumer price changes are reflected contemporaneously in house prices in depressed market conditions and the full effect of consumer price changes is contemporaneously impounded in house prices in buoyant conditions. Nevertheless, our estimates indicate that adjustment to equilibrium is characterised by asymmetries and some degree of overshooting. Thus the housing market, while being "moderately efficient", retains the capability for delivering surprising and potentially destabilising episodes.

**Table 1: Sales price summary statistics, population and activity**

| Regional council | 2002 median sales price ($000)* | Real % price change: 1981–2002 | Average no. quarterly sales | Population (2001 census) | Population density (2001 census) | % change in real economic activity: 1981–2002 |
|---|---|---|---|---|---|---|
| RC01 | 157 | 46 | 568 | 140,133 | 10.1 | 105 |
| RC02 | 282 | 111 | 5349 | 1,158,891 | 206.9 | 98 |
| RC03 | 166 | 61 | 1658 | 357,726 | 14.0 | 92 |
| RC04 | 168 | 38 | 1270 | 239,412 | 19.2 | 84 |
| RC05 | 100 | 2 | 168 | 43,974 | 5.3 | 32 |
| RC06 | 142 | 32 | 616 | 142,947 | 10.1 | 64 |
| RC07 | 106 | 12 | 524 | 102,858 | 14.1 | 73 |
| RC08 | 98 | 12 | 1191 | 220,089 | 9.9 | 51 |
| RC09 | 203 | 87 | 2206 | 423,765 | 52.2 | 77 |
| RC11 | 162 | 46 | 669 | 122,475 | 5.4 | 97 |
| RC12 | 63 | 24 | 173 | 30,303 | 1.3 | 62 |
| RC13 | 146 | 58 | 2727 | 481,431 | 10.6 | 99 |
| RC14 | 117 | 38 | 1158 | 181,542 | 5.7 | 59 |
| RC15 | 66 | –27 | 578 | 91,002 | 2.6 | 46 |

*Expressed in $NZ. On average, over 2002, NZ$1 = US$0.46.

**Table 2: Results of panel unit root tests**

| | Levin-Lin-Chu | | Im-Pesaran-Shin | |
|---|---|---|---|---|
| **Variable** | **Trend and constant** | **Constant** | **Trend and constant** | **Constant** |
| $\mathbf{P_z}$ | I(0)/I(1) | I(0)/I(1) | I(0)/I(1) | I(0)/I(1) |
| $\mathbf{Y_z}$ | I(1) | I(1) | I(1) | I(1) |
| $\mathbf{S_z}$ | I(0)/I(1) | I(0)/I(1) | I(0) | I(0) |
| $\mathbf{H_z}$ | I(0) | I(0) | I(0) | I(1) |
| $\mathbf{UC2_z}$ | I(1) | I(0) | I(1) | I(0) |
| $\mathbf{UC3_z}$ | I(1) | I(1) | I(1) | I(1) |
| $\mathbf{COMP_z}$ | I(0) | I(0) | I(0) | I(0) |

**Table 3: Long-run house price estimates**

| | UC1 | | UC2 | | UC3 | | UC4 | |
|---|---|---|---|---|---|---|---|---|
| | **OLS** | **SUR** | **OLS** | **SUR** | **OLS** | **SUR** | **OLS** | **SUR** |
| $\mathbf{Y_{zt}}$ | 1.372<br>(24.87) | 1.198<br>(24.43) | 1.169<br>(22.82) | 0.921<br>(18.49) | 1.575<br>(30.42) | 1.331<br>(27.02) | 1.577<br>(29.89) | 1.338<br>(26.75) |
| $\mathbf{H_{zt}}$ | -0.582<br>(7.30) | -0.598<br>(10.93) | -0.691<br>(9.43) | -0.646<br>(11.89) | -0.671<br>(7.90) | -0.723<br>(12.46) | -0.470<br>(5.95) | -0.456<br>(8.37) |
| $\mathbf{UC_{zt}}$ | -0.0057<br>(10.46) | -0.0062<br>(9.26) | -0.0077<br>(20.06) | -0.0097<br>(24.66) | -0.0044<br>(7.92) | -0.0053<br>(8.14) | -0.0031<br>(3.16) | -0.0038<br>(2.85) |
| $\mathbf{UCD_t}$ | 0.0759<br>(7.52) | 0.0625<br>(4.85) | 0.0500<br>(5.96) | 0.0223<br>(1.87) | 0.1154<br>(13.55) | 0.1013<br>(8.40) | 0.1093<br>(8.58) | 0.0879<br>(5.04) |
| **Adj. $R^2$** | 0.9979 | 0.9729 | 0.9980 | 0.9780 | 0.9974 | 0.9724 | 0.9979 | 0.9706 |
| **s.e.** | 0.0542 | 0.0545 | 0.0485 | 0.0491 | 0.0544 | 0.0550 | 0.0560 | 0.0567 |
| **Obs** | 1,232 | 1,232 | 1,232 | 1,232 | 1,232 | 1,232 | 1,232 | 1,232 |

All equations are estimated over 1981(1)–2002(4). Absolute t statistics in parentheses.
Each equation has unrestricted constants, TIME and $TIME^2$, included but not reported.
The dependent variable is the log of real house prices ($P_{zt}$), where z denotes region,and t denotes time.
$Y_{zt}$ is the log of real regional economic activity.
$H_{zt}$ is the log of housing stock.
$UC_{zt}$ is the user cost of capital [definitions as in the text, entered only from 1985(1) onwards].
$UCD_t$ is a dummy variable equal to 1 to 1984(4) and equal to 0 thereafter to account for the regulated financial system.

**Table 4: Dynamic house price estimates**

| | (1) OLS symmetric | (2) SUR symmetric | (3) PW symmetric | (4) GLS symmetric | (5) OLS asymmetric | |
|---|---|---|---|---|---|---|
| | | | | | -ve $RES_{t-1}$ | +ve $RES_{t-1}$ |
| **$RES_{zt-1}$** | -0.4326*** (18.73) | -0.4255*** (18.66) | -0.4681*** (16.22) | -0.4507*** (19.64) | -0.7473*** (27.34) | -0.7196*** (26.41) |
| **$\Delta Y_{zt}$** | 0.6320*** (7.16) | 0.5682*** (6.43) | 0.5925*** (6.45) | 0.5211*** (6.73) | 0.9293*** (9.94) | 0.8949*** (9.77) |
| **$\Delta H_{zt}$** | -0.7900*** (2.70) | -0.6266** (2.14) | -0.7202** (2.43) | -0.6298** (2.38) | 0.2383 (0.79) | -0.8109*** (2.66) |
| **$\Delta UC2_{zt}$** | -0.0017** (2.36) | -0.0017** (2.45) | -0.0016** (2.20) | -0.0017*** (2.86) | -0.0057*** (6.83) | -0.0056*** (8.04) |
| **$COMP_{zt}$** | 0.1490*** (8.40) | 0.1436*** (8.11) | 0.1673*** (8.02) | 0.1464*** (8.59) | 0.1184*** (5.87) | 0.0827*** (4.85) |
| **$S_{zt-2}$** | 1.2656*** (5.78) | 1.2765*** (5.83) | 1.1701*** (4.84) | 1.0693*** (5.31) | 1.3476*** (5.85) | 0.0652 (0.29) |
| **$\Delta PC_t$** | -0.4941*** (6.02) | -0.4787*** (5.82) | -0.5366*** (6.29) | -0.5409*** (7.53) | -0.2645*** (3.28) | 0.0138 (0.14) |
| **Obs.** | 1204 | 1204 | 1204 | 1204 | 1204 | |
| **Adj. $R^2$** | 0.3057 | 0.3046 | 0.3251 | | 0.6337 | |
| **s.e.** | 0.0366 | 0.0366 | 0.0361 | | 0.0266 | |

The dependent variable is the log change in real house prices ($\Delta P_{zt}$), where z denotes region, t denotes time.
$RES_{zt-1}$ is the lagged long-run house price residual from Table 3.
$\Delta Y_{zt}$ is the log change in real regional economic activity.
$\Delta H_{zt}$ is the log change in housing stock.
$UC2_{zt}$ is change in user cost of capital (the second proxy for $\dot{g}$ in the text).
$COMP_{zt}$ is a housing stock composition variable.
$S_{zt-2}$ is house sales to housing stock ratio lagged two quarters.
$\Delta PC_t$ is the log change in consumer prices.
An unrestricted constant term is included but not reported.
Absolute t statistics in parentheses; * indicates significant at 10%; ** significant at 5%; *** significant at 1%.

**Table 5: Sales ($S_{zt}$) response to house price developments**

| | **OLS** |
|---|---|
| $RES_{zt-1}$ | –0.0196<br>(–7.70)*** |
| $\Delta P_{zt}$ | 0.0080<br>(3.34)*** |
| $\Delta P_{zt-1}$ | 0.0258<br>(10.01)*** |
| $\Delta P_{zt-2}$ | 0.0186<br>(7.89)*** |
| $\Delta P_{zt-3}$ | 0.0081<br>(3.60)*** |
| $\Delta P_{zt-4}$ | 0.0073<br>(3.30)*** |
| $\Delta P_{zt-5}$ | 0.0059<br>(2.76)*** |
| Adjusted $R^2$ | 0.69 |

The dependent variable is house sales to housing stock ratio ($S_{zt}$) where z denotes region, t denotes time.
$RES_{z-1}$ is the lagged long-run house price residual from Table 3 column (1).
$\Delta P_{zt}$ is the log change in real house price.
Unrestricted constants TIME and $TIME^2$ included but not reported.
Absolute t statistics in parentheses; *** indicates significant at 1%.

**Figure 1: Real house price response to simulated changes in economic activity**

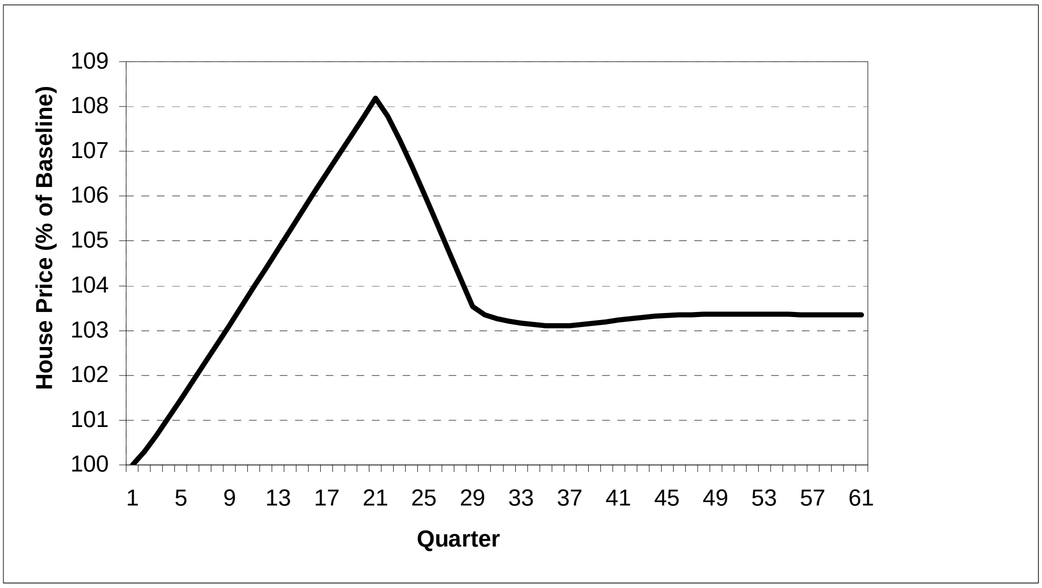
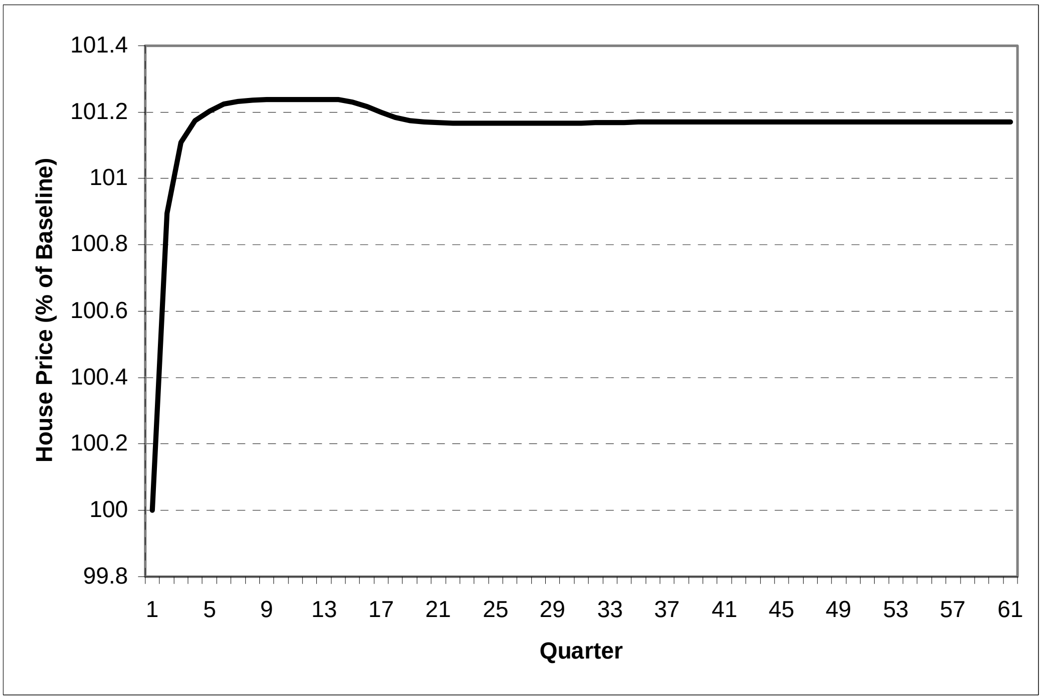



**Figure 2: Real house price response to 1% increase in economic activity**



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