

## The Response of Australian Consumption to Housing Wealth

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### **Abstract**

Large variations in house prices can lead to significant changes in the level of household wealth and this may affect household consumption. Using Lettau and Ludvigson's (2004) cointegration approach, we investigate the response of non-housing consumption to permanent and transitory changes in financial and non-financial (housing) wealth in Australia since the mid-1970s. The data provide evidence that housing wealth contains a large transitory component, but up to 2004 at least, these transitory changes in wealth are not associated with any significant response in consumption. When post-2004 data are included in the estimation, there is some evidence that household consumption responds to recent transitory rises in wealth and labor income. However this finding needs to be weighed against weaker evidence of a cointegrating relationship for consumption, income and wealth.

**Keywords:** consumption, wealth, housing wealth, cointegration, permanent and transitory shocks

**JEL Classification:** E21, C32

## 1. Introduction

A longstanding empirical question in macroeconomics is the relationship between household consumption and wealth. A key issue, particularly for macroeconomic policy-makers, is the extent to which changes in wealth, due to large variations in asset prices, influence the level of household consumption and subsequently the level of aggregate demand. Traditionally researchers have attempted to answer this question by estimating the marginal propensity to consume (MPC) out of aggregate wealth or its components (see Ludvigson and Steindel, 1999; Poterba, 2000; Tan and Voss 2003; Case, Quigley and Shiller, 2005). Estimates of the MPC for wealth are frequently obtained using single equation methods, often under the assumption that changes in wealth can be treated as exogenous. Implicit in such an approach is the assumption that all of the adjustment to a change in wealth must come through changes in consumption.

In two recent papers, Lettau and Ludvigson (2001, 2004) demonstrate that estimates of wealth effects obtained by conventional regression methods can be very misleading. Instead they argue that in general, understanding the relationship between consumption and wealth requires a systems approach, where both consumption and wealth are allowed to behave endogenously.<sup>1</sup> To illustrate this point Lettau and Ludvigson (2004) show that for the United States (U.S.) much of the post-war variation in household wealth, although highly persistent, is transitory in nature eliciting little or no response in household consumption.<sup>2</sup> As a result, large movements in wealth are found to have only a small impact on consumption, much smaller than the estimates implied by earlier U.S. studies. The problem with conventional estimates of wealth effects is that

they can overstate the adjustment of consumption to changes in asset wealth because any future adjustment in wealth itself is ignored.

One feature of the U.S. wealth data that emerges from Lettau and Ludvigson's work is the importance of fluctuations in stock prices in driving the transitory component of wealth.<sup>3</sup> The stock market boom of the 1990s and its subsequent correction is a prominent example. However recent events such as the U.S. house price boom and bust have re-focused policy-makers attention from the effects of stock prices on wealth to the effects of fluctuations in house prices. While national housing booms are a somewhat recent feature of the data for many economies, Australia is one country that has exhibited a reasonably pronounced housing cycle over a number of years.<sup>4</sup> Moreover in the Australian economy the private housing stock represents around sixty percent of household wealth, implying that fluctuations in house prices can have a large effect on total household wealth.

In this paper, we adopt the methods of Lettau and Ludvigson (2004) to investigate the response of household consumption to permanent and transitory fluctuations in financial and non-financial (housing) wealth in Australia over the period 1976–2008.<sup>5</sup> The first task is to identify a stable long-run relationship between consumption, labor income and the components of household wealth. Such a relationship is evident in the data but only up until 2004, after which it appears to break down due to the unprecedented rapid rises in household wealth and labor income associated with the commodity boom in Australia commencing in 2004. Nonetheless, we are able to use the information from the long-run analysis to assess the dynamic relationship between consumption and household wealth. Our principal conclusion is that for the greater part

of our sample, transitory changes in wealth are not associated with any significant response in consumption; however, when we consider the most recent data, the evidence is much less clear-cut.

## 2. Theory

### *The Household Intertemporal Budget Constraint*

We begin by generalizing Lettau and Ludvigson's (2001, 2004) theoretical framework, decomposing non-human wealth  $A$  into two sub-components; financial wealth  $A_{ft}$  and non-financial wealth  $A_{nt}$  i.e.  $W_t = H_t + A_{ft} + A_{nt}$  and  $H_t$  is human wealth. As shown in the Appendix the log-linear approximation to the household's intertemporal budget constraint is now given by

$$c_t - \beta_f a_{ft} - \beta_n a_{nt} - \beta_y y_t \approx E_t \sum_{j=1}^{\infty} \rho^j (\beta_y \Delta y_{t+j} + \beta_f r_{aft+j} + \beta_n r_{ant+j} - \Delta c_{t+j}) \quad (1)$$

where  $r_{aft}$  and  $r_{ant}$  are the real returns to financial and non-financial wealth respectively. The coefficients  $\beta_f$ ,  $\beta_n$  and  $\beta_y$  are the steady-state shares  $A_f/W$ ,  $A_n/W$  and  $H/W$  and should satisfy the restriction  $\beta_f + \beta_n + \beta_y = 1$ . The left-hand side of equation (1) is a generalization of Lettau and Ludvigson's  $ca$  variable to  $ca_f a_n y$ . Provided the right-hand side variables are stationary, equation (1) implies the existence of a cointegrating relationship for consumption, labor income, financial wealth and non-financial wealth.

Since equation (1) is derived from the household budget constraint it is an (approximate) identity. To see this consider a shock that causes the left-hand side variable  $ca_f a_n y$  in equation (1) to become positive (this corresponds to dis-saving by the

household). For the intertemporal budget constraint to hold it is necessary that some combination of the following changes occur. Either labor income growth or real returns on financial or non-financial wealth are expected to increase in the future or consumption growth is expected to decline. By itself the intertemporal budget constraint makes no prediction about which of these changes will occur in order to maintain intertemporal budget balance.

To convert the budget constraint (1) into a behavioural model for consumption it is necessary to make some assumption about the determinants of expected consumption growth. A simple assumption typically associated with the permanent income hypothesis is that the log of consumption is approximately a random walk (possibly with drift).<sup>6</sup> In this case equation (1) becomes,

$$c_t - \beta_f a_{ft} - \beta_n a_{nt} - \beta_y y_t \approx E_t \sum_{j=1}^{\infty} \rho^j (\beta_y \Delta y_{t+j} + \beta_f r_{aft+j} + \beta_n r_{ant+j}) \quad (2)$$

which can be interpreted as a generalization of Campbell's (1987) "saving for a rainy day" formulation of the permanent income hypothesis, to allow for time-varying asset returns. An attractive feature of equation (2) is that it treats the expected real returns to financial and non-financial assets in a similar manner to expected changes in labor income. Provided these real return series are stationary then we can see from equation (2) that the immediate effect of anticipated future changes in asset returns is on the saving variable  $ca_f a_n y$  rather than on the level of consumption.

### *A Long Run Model*

Equation (1) forms the basis for our empirical analysis of the Australian consumption and wealth data. We can re-write equation (1) as

$$c_t - \beta_f a_{ft} - \beta_n a_{nt} - \beta_y y_t = \beta_0 + \eta_t \quad (3)$$

where the slope coefficients are steady-state share of wealth parameters and  $\beta_f + \beta_n + \beta_y = 1$ . The constant term arises from the linearization of the budget constraint. The error term  $\eta_t$  is assumed to be mean zero stationary random variable. It consists of expected future log-differences of consumption and labor income (discounted) as well as expected future net returns on the different asset components (also discounted). Provided the variables on the right-hand side of equation (1) are stationary then the linear combination on the left-hand side must form a stationary or cointegrated vector. Assuming that  $c_t$ ,  $y_t$ ,  $a_{ft}$ , and  $a_{nt}$  are first-difference stationary and that  $r_{aft}$  and  $r_{ant}$  are stationary in levels, equation (3) represents a cointegrating relationship.

Re-writing equation (3) as

$$c_t = \beta_0 + \beta_f a_{ft} + \beta_n a_{nt} + \beta_y y_t + \eta_t \quad (4)$$

we obtain what looks like a standard formulation for the long-run consumption function, *albeit* in logarithms rather than levels. However it is a straightforward exercise to convert the elasticities in (4) into long-run marginal propensities for wealth and labor income. Now Lettau and Ludvigson (2004) point out that if one interprets (4) as a long-run consumption function and uses it to estimate the effect of wealth on consumption, then what  $\beta_f$  (and  $\beta_n$ ) measures is the *ceteris paribus* effect of a permanent change in wealth on consumption. If all or even most changes in wealth are permanent this would be all

that is required. However if wealth is also affected by quantitatively important transitory shocks then it may be inappropriate to use  $\beta_f$  or  $\beta_n$  to predict the effect on consumption of some anticipated change in wealth. If households view a given change in wealth as transitory then using these long-run parameters to estimate the MPC will overstate the predicted change in consumption. The relative importance of transitory and permanent shocks to wealth cannot be inferred from long-run models like (4). To estimate the response of consumption to a transitory wealth shock requires estimation of the vector error-correction (VEC) model associated with (4).

### *Is Housing Really Wealth?*

In generalizing Lettau and Ludvigson's model we have implicitly assumed that an agent's housing wealth can be treated in a manner that is analogous to their financial wealth. While this is a standard assumption in the empirical literature on consumption, its validity has been questioned by a number of authors who argue there is a fundamental difference between financial and housing wealth (Aoki, Proudman and Vlieghe, 2004; Carroll, 2004 and Buiter, 2008a,b). The basic idea is straightforward. For the average home-owner, a rise in house prices has two effects. It increases the value of their asset; however it also raises the opportunity cost of consuming the flow of services derived from the house. Consider a simple one-period budget constraint,

$$P_{nh}C_{nh} + P_hC_h = A_f + P_hH^* \quad (5)$$

where  $C_{nh}$  is non-housing consumption,  $C_h$  is housing consumption,  $A_f$  is nominal financial wealth and  $H^*$  is the housing endowment. The price per unit of non-housing consumption is  $P_{nh}$  and the price per unit of housing services is  $P_h$ . House prices enter

both sides of the budget constraint and their net effect depends on the difference  $(H^* - C_h)$ . Where an agent is an owner-occupier and just consumes their endowment, housing wealth drops out of the budget constraint entirely. In effect any contribution to wealth from the increase in house prices is offset by the rise in the opportunity cost of housing services. Buitert (2008a,b) shows that the absence of a wealth effect from changes in house prices also carries over to general dynamic models. He establishes the lack of a wealth effect on consumption from changes in house prices (provided they are driven by economic fundamentals) in a standard representative agent model and in the Yarrri-Blanchard OLG model (Yarri, 1965; Blanchard, 1985).

While there is debate about the existence of a traditional wealth effect from housing, there are other mechanisms by which changes in housing wealth may affect consumption. These mechanisms relate to financial innovation and product development in the provision of housing finance. For Australia these developments began in the 1990s following the deregulation of financial markets in the 1980s. During the early 1990s, specialist mortgage originators entered the mortgage market. Given the high margins at the time, housing loans were extremely profitable and the new entrants competed aggressively by offering a broad range of products and standard variable mortgage rates considerably below those offered by established lenders. Established lenders responded to this competition by offering enhanced lending products and by reducing their margins as well. Important among these new lending products were home equity loans which allowed owner-occupiers to borrow using home equity as collateral. In particular, these loans permitted borrowers to refinance debt at lower interest rates made possible by the collateral backing of home equity. The use of these loans has increased with the rise in

the price of houses. Since 2003 the stock of debt secured by housing equity as collateral has increased by more than household spending on new housing, renovations and housing transfer costs so that the household sector has extracted equity from the housing stock (Schwartz, Hampton, Lewis and Norman, 2008). By contrast, in the decades prior to 2003, the household sector consistently injected equity into the housing stock by paying down their mortgage principal or by using some of their own funds to finance renovations. Financial innovation in the mortgage market has also allowed first-time home buyers to secure loans for which the ratio of the loan to their personal income (LTY) or to the value of their home (LTV) is substantially higher than had been the case previously.

At issue is whether these developments have had an effect on household consumption. Muellbauer (2007) has argued that the higher LTV and LTY ratios have meant that first-time home buyers who are typically the young have not needed to save as much as previously for a house deposit in a rising housing market. This has diminished the negative impact of rising house prices on the consumption of the young. At the same time, older households who are mainly owner-occupiers have an enhanced capacity to increase their consumption as house prices rise through increased housing collateral. The expected net impact is for consumption to rise with rising house prices. Some evidence of this is found by Muellbauer in U.K and U.S. data. For Australia, Schwartz et al (2008) report the results of a comprehensive survey, commissioned by the Reserve Bank of Australia, of the house equity withdrawal and injections decisions of Australian households in 2004. The survey found that around two-thirds of the equity withdrawn in 2004 was invested in other assets or used to pay down other loans. In contrast, only a

relatively small proportion of equity withdrawn (18 percent) was used to fund consumption which suggests that there is not a close link between consumption and housing equity withdrawal in Australia.

In a theoretical framework, Aoki, Proudman and Vlieghe (2004) emphasize the collateral channel. To the extent that households require collateral to borrow, rising house prices will increase housing equity and provide the opportunity to borrow at lower interest rates. In this case consumption will tend to respond to changes in house prices and housing wealth in their model. However Buitier (2008b) makes the point that this collateral effect does not change the present-value of a household's consumption stream, but will tend to raise current consumption and reduce future consumption – a transitory effect. Aggregate effects on consumption can arise from the distributional effects of house price changes if households have different marginal propensities to consume.

Finally Buitier (2008b) notes that if a change in house prices is due to a rational speculative bubble rather than being driven by economic fundamentals, then there is a pure wealth effect on consumption. During a bubble the price of housing rises, but there is no corresponding rise in the present-value of fundamental housing services. Owner-occupiers will experience a real wealth effect and can be expected to respond via higher non-housing consumption. If a house price bubble was viewed as a permanent feature of the housing market, it would enter equation (4) presumably as a component of measured housing wealth. Where bubbles are stochastic and bust with some probability, their effects are likely to show up in the transitory component of non-financial wealth, in which case rational consumers would respond with relatively small changes in consumption.

Inspection of equation (5) suggests that if the value of housing services is correctly measured, then aggregate consumption will move approximately one-for-one with changes in housing wealth. However there is relatively little economic interest in such a finding. Rather in considering whether there are direct or indirect wealth effects associated with housing, a more sensible strategy seems to be to focus on the behavior of non-housing consumption expenditure. If there is no pure wealth effect from housing and no important collateral channel then we would expect to see little response in non-housing consumption to changes in housing wealth. Furthermore we would argue that for policy-makers the important issue is whether housing wealth has effects on non-housing consumption, rather than on housing consumption (a large part of which is imputed).

### **3. Empirical Results**

#### *Data*

Consumption is measured by total household final consumption expenditure less household final consumption expenditure on rent and other dwelling services. The latter comprises rent paid by householders to the owners of dwellings, the imputed value of housing services accruing to owner-occupiers and expenditure on other dwelling services, for example, local council rates and water and heating. Expenditure on rent and other dwelling services increased as a proportion of total household final consumption expenditure from 14.6 percent in December 1976 to 19 percent in September 1993. Thereafter, it declined slowly to 17.5 percent in September 2008.

Labor income is the income derived by households from the supply of labor net of tax. It includes transfer payments, which are predominantly social benefit payments. Financial wealth is the total value of financial assets held by households net of financial

assets owing. Non-financial wealth consists of the value of dwelling assets and consumer durables, net of debt. The precise definition and construction of the labor income and wealth series are described in Tan and Voss (2003). The series are divided by the implicit price deflator for total household final consumption expenditure excluding expenditure on rent and other dwelling services. This implicit price deflator is used because it corresponds to our consumption measure and is therefore consistent with the intertemporal budget constraint following the reasoning of Rudd and Whelan (2006). In one respect, however, our series are not consistent with the intertemporal budget constraint since the value of durables should be excluded from non-financial wealth given that expenditure on durables is included in our consumption measure. This measurement error, though unlikely to be important, was unavoidable because the series for non-financial wealth net of consumer durables is only published by the Reserve Bank of Australia from June 1988, well after the commencement of our sample. The real series are divided by the resident population to obtain real per-capita series, the logarithms of which are used in the empirical analysis. Finally, the series are quarterly and the full sample is 1976:4 – 2008:3.

*Is there a stable long-run relationship?*

The derivation of equation (1) is based on the assumption that the shares of human, financial and non-financial wealth in total wealth are constant in the steady-state. Carroll, Otsuka, and Slacalek (2006) argue this is a strong assumption and provide theoretical reasons why there are likely to be changes in the steady-state wealth shares. A changing economic environment is one reason, and it is the case that like many countries Australia has undergone significant structural changes to its financial and labor markets

over the past three decades. Nevertheless in practice the stability of the long-run coefficients in (1) seems to be an empirical issue and consequently we consider a number of different types of evidence.

Figure 1 presents our measures of the three wealth shares. Since labor income is the return on human capital, we capitalize it using  $Y(t)/0.025$ , where we are assuming a ten percent annual return to human capital.<sup>7</sup> The shares are not constant and appear to exhibit trends. The share of human capital wealth has declined steadily over the sample while the share of non-financial and financial wealth has risen. For non-financial wealth, which is primarily housing, three episodes of strong housing price growth are evident in the late 1980s, the late 1990s, and the early 2000s. Since 2004, the share of non-financial wealth in Australia has stabilized as house price growth has stabilized over this period. For financial wealth, episodes of strong equity price growth are evident in the late 1980s, the late 1990s and most recently the years 2003–06; the latter is associated with the spectacular growth in the value of resource stocks as a consequence of China's unprecedented demand for resources. This led to substantial improvement in the terms of trade and to very strong growth in labor income. Over the last two years, 2007–08, with the advent of the financial crisis, these strong gains have started to reverse themselves and this is evident in the declining asset shares. As global economic conditions continue to deteriorate through 2009, one expects the share of financial and non-financial wealth to continue downward. Formally, a test of the null hypothesis that each of the wealth shares has a unit root cannot be rejected using either the Phillips-Perron (PP) or the Augmented Dickey-Fuller (ADF) tests, further identifying a potential problem with parameter instability in the long-run relation.

In light of figure 1 we begin our empirical analysis with an examination of the stability of the cointegrating relationship predicted by equation (1). Figure 2 shows the results from the PP unit root test (Phillips and Perron, 1988) applied to the residuals from an OLS regression of consumption on a constant, labor income, financial and non-financial wealth.<sup>8</sup> The initial sample over which the OLS regression is estimated is 1976:4 – 1990:1. The sample is then extended by one quarter at a time to 2008:3. For each sample, the PP statistic is calculated from the OLS regression residuals where the autocovariance function is truncated at four lags. The PP statistic provides a test of the null hypothesis of no cointegration. The 10 percent critical value of the test statistic ( $-3.84$ ) is shown by the horizontal line. The figure shows that with the exception of a possible break in the long-run relationship in the sample that ends in 1999:1 (shown by the spike above the horizontal line) there is evidence for a stable long-run relationship in samples to 2004:3. The PP statistic then rises dramatically so that for the samples that end from 2006:3 there is no evidence for cointegration.<sup>9</sup> The figure suggests that the long-run relationship starts to break down near 2004:3.

We also considered a version of the *SupF* test proposed by Hansen (1992) to test for parameter stability in the cointegrating regression.<sup>10</sup> The null hypothesis of the test is that all the coefficients in the cointegrating relation including the constant are stable and the alternative is a single structural break of unknown date in at least one of these coefficients. The test entails computing a sequence of Chow-tests for structural change in the cointegrating regression conducted for a set of possible break dates through the middle 70 percent of the sample (1981:2–2004:3). The largest F-statistic (*SupF*) is 208.8, far larger than the 5 percent critical value tabulated by Hansen, and it occurred at 1998:4.

The *SupF* and PP statistics are in broad agreement that there was a break in the cointegrating relation around this time. The *SupF* test, however, does not provide us with any information beyond the last break point considered, 2003:4 so we must rely on the information from the recursive PP statistics for the final part of the sample.

The recursive PP statistics indicate a changed structural relationship from 2004:3 and possibly a break earlier in the sample which was supported by the *SupF* statistic. Based on these findings we report in table 1 the results of tests for cointegration and the parameter estimates for the full sample (1976:4-2008:3) and for the sub-samples to 2004:3 and 1998:4, respectively. Part A of the table reports the results of the PP test (shown in figure 2), the  $Z_\tau$  test (Phillips and Ouliaris, 1990) and the ADF test (Dickey and Fuller, 1981). Like the PP test, the  $Z_\tau$  and ADF tests have a null of no cointegration and test for a unit root in the residuals from an OLS regression of consumption on a constant, labor income and financial and non-financial wealth (Engle and Granger, 1987).<sup>11</sup> For the full sample, all three tests do not reject the null of no cointegration. However, both the PP and  $Z_\tau$  tests reject the null of no cointegration at the 5 percent level for the sample to 2004:3 and reject at the 10 percent level for the sample to 1998:4. The ADF test does not reject the null in both sub-samples but the rejection is only marginally above the 10 percent level for the sample to 2004:3.<sup>12</sup> Although there is no evidence for a cointegrating relationship in the full sample, this appears to be a recent phenomenon. For the sub-samples where we do find evidence of cointegration, it is worth noting that there is relatively strong evidence for cointegration, particularly for the sample ending in 2004:3, and the similarity of the estimated coefficients seems to belie the concerns raised about the non-constant share parameters, which tend to trend over the

sub-samples as well as the full sample. We view this as support for using the long-run model as a guide for our empirical work – at least for these sub-samples.

Part B of table 1 presents the dynamic ordinary least squares (DOLS) estimates of the coefficients in the cointegrating relation. This method of estimation is due to Stock and Watson (1993) and it provides efficient estimates of the parameters in a cointegrating relation. The DOLS regression augments the OLS regression with  $k$  leads and lags of the first difference of the right-hand side variables.<sup>13</sup> Since the choice of  $k$  is somewhat arbitrary, a range of values from 2 to 4 was tried. The table reports the parameter estimates for  $k=4$ , however the parameter estimates are broadly similar for  $k=2$  and  $k=3$ . The log-linear budget constraint predicts that the estimated coefficient on labor income and the components of wealth should sum to unity as these coefficients represent the steady-state shares of human wealth and the components of non-human wealth in total wealth, respectively. However, this prediction is not quite accurate in the current context because consumption excludes expenditure on housing services. Specifically, if it is assumed that the log of total household consumption is proportional to the log of non-housing consumption (the factor of proportionality  $\lambda$  is greater than one), the theory would predict that the coefficients in the cointegrating relation sum to  $1/\lambda$ , a number slightly less than one.<sup>14</sup> The DOLS estimates are broadly consistent with this prediction. They sum to slightly larger than one in the samples for which strong evidence of cointegration is found (1.07 and 1.01 for the samples ending in 2004:3 and 1998:3, respectively) and sum to 0.925 for the full sample. The estimated coefficients are remarkably similar where strong evidence for cointegration is found. The estimated coefficient on financial wealth is small, around 0.03, while the estimated coefficient on

non-financial wealth is large, around 0.21, in both sub-samples. Also, the first-order autocorrelation coefficient of the DOLS residuals are both considerably less than one, consistent with the finding of cointegration. In the full sample where there is no evidence of cointegration, the coefficient on financial wealth is now three times as large, around 0.10, while the coefficient on non-financial wealth of around 0.13 is considerably smaller.<sup>15</sup> From 2004-2007, the Australian stock price index doubled resulting in a rapid increase in financial wealth and inclusion of this data causes the estimated coefficient on financial wealth to triple in size and the coefficient on non-financial wealth to fall substantially in the DOLS regression.<sup>16</sup>

Figure 3 provides a graph of the residual  $ca_f a_n y$  to the end of the sample constructed from the DOLS estimates for the full sample and for the sub-sample to 2004:3, respectively. Towards the end of the sample there is an extended sequence of very large negative values for both residuals, which raises questions about the stability of the model. A natural question to ask is what might be driving these sequences of residuals. To determine this, consider the decomposition of the predicted values from the 1976:4-2004:3 model presented in table 2. The decomposition is done in annual growth rates for the whole of the sample as well as the period 2002:1 to 2004:3. This latter period is where we observe the sequence of negative residuals. We focus on the 1976-2004 sample since we have evidence of cointegration for this sample. From table 2, there are two obvious sources underlying the negative residuals: the strong growth over this period in non-financial asset values (housing) and the strong growth in labor income. While financial asset growth was significant over this period, it was much smaller than non-financial assets and, moreover, receives a much smaller weight in the long-run

relationship. A similar exercise for the 2008 sample provides similar conclusions though in this case the labor income contribution is even more dominant.

This recent behavior of  $ca_f a_n y$  suggests two interpretations of the data. The first is that the estimated coefficients in  $ca_f a_n y$  are not stable so that the relationship between consumption, labor income and the two components of wealth is not characterized by the cointegration framework adopted here. The other interpretation is that there is a stable cointegrating relationship and that this relationship will be re-established in future data on the expectation that this data will accord more with historical experience than the most recent data in our sample. In that case, the estimated parameters will more closely resemble those estimated in the samples for which strong evidence for cointegration was found. Implicit in this view is the requirement that there will be some future adjustment in the series which returns  $ca_f a_n y$  to zero, thereby restoring the long-run relationship. To determine how the series adjust requires estimation of the vector-error correction (VEC) model to which we now turn.

#### *Vector Error Correction Models*

The VEC model is given by,

$$\Delta X_t = \mu + \Pi X_{t-1} + \Pi_1 \Delta X_{t-1} + \dots + \Pi_{k-1} \Delta X_{t-k+1} + \varepsilon_t, \quad (6)$$

where  $X_t$  is the vector of the series,  $\Delta$  is the first difference operator,  $\mu$  is a vector of constant terms,  $\Pi_i$  is the matrix of coefficients on the  $i$ 'th lagged change in  $X_t$  and  $\varepsilon_t$  is a vector of serially uncorrelated random disturbances with mean zero and covariance matrix  $\Omega$ . Under cointegration,  $\Pi = \alpha\beta'$  where  $\beta$  is the column vector of coefficients on

the series in the cointegrating relation and  $\alpha$  is the column vector of adjustment coefficients on the error-correction term  $\beta' X_{t-1}$  in each equation. Each equation in the VEC model is estimated by OLS, conditional on the DOLS estimates of the parameters in the cointegrating relation. Importantly, the adjustment coefficient on the error-correction term in each equation indicates how each variable in the system adjusts to restore long-run equilibrium following a shock to the error correction mechanism. Thus an understanding of the adjustment process to restore long-run equilibrium among the series requires the estimation of the vector error correction system.

Table 3 reports the estimates of the VEC model for the sub-sample that ends in 2004:3. We report these results since there is strong evidence for cointegration in this sample so that an error-correction mechanism is operating. There is also strong evidence for cointegration in the sample to 1998:4 but we do not report the results from this sample as they are very similar to those from the 2004 sample. For this model, the AIC criterion selected three lags. The estimated coefficient on the error-correction term in the non-financial wealth equation is large, positive and statistically significant at the 5 percent level. In the consumption equation, this coefficient is negative and significant and is three times smaller (in absolute value) than the coefficient on non-financial wealth. The estimated error-correction coefficient is not significant in either the labor income or financial wealth equation. These results imply that when private saving is low (the error-correction term is positive), consumption is predicted to fall in the future and non-financial wealth is predicted to rise thereby restoring the long-run relationship while labor income and financial wealth do not directly enter into the adjustment process. Turning to the short-run dynamics, the sum of the estimated coefficients on lagged changes in non-

financial wealth is statistically significant in the consumption, labor income and non-financial equations. Lagged changes in financial wealth also predict current consumption growth. These findings for consumption formally reject the martingale form of the permanent income hypothesis (Tan and Voss, 2003). There does not appear to be any significant diagnostic problems apart from non-normality and heteroskedasticity in the non-financial wealth residuals. These appear due to outliers (excess kurtosis) rather than to skewness in the distribution of the residuals so that reliable inference can still be made.

The estimated VEC equation for consumption shows that changes in wealth, particularly non-financial wealth, have predictive power for consumption growth next period, consistent with the findings of single equation studies (Muellbauer, 2007). Single equation models for consumption typically include many variables in the estimation, for example, consumer sentiment, credit conditions, and debt-servicing capacity. The OLS estimates of the coefficients in the VEC consumption growth equation are consistent. Lettau, Ludvigson and Barczi (2001) show that inclusion of such additional explanatory variables will lead to more efficient estimates of the coefficients provided these variables are in fact important short-run determinants of consumption growth. While there is an efficiency gain in finite samples, the estimation must be performed by instrumental variables when, as seems likely, the additional explanatory variables are not orthogonal to the cointegrating vector. Only then can the estimated coefficient on the error-correction term be reliably interpreted as the response of consumption growth to variation in the disequilibrium error independent of only past movements in the variables in the cointegrating relation itself, consistent with the long-run model. The VEC framework has the advantage that it can uncover predictability in the growth of the series over long

horizons. Because the error-correction term is significant in the consumption equation, there is a measure of predictability in consumption growth over long-horizons implying that consumption adapts to permanent innovations in wealth and income. However, consumption's role in the adjustment process is less important than non-financial wealth because the coefficient on the error-correction term in its equation is considerably less, by more than one third in absolute value, than the coefficient in the non-financial wealth equation so that consumption reverts far more quickly to trend.

Table 3 also reports the results from estimating the full sample VEC model for which the AIC criterion also selected three lags. These results should be interpreted with some caution as the test statistics reported in table 1 do not provide strong evidence against the null of no cointegration for the full sample. However this may reflect the low power of these tests and stronger evidence of cointegration is obtained from the estimates of the VEC model (Zivot, 2000). The estimated coefficient on the error-correction term is positive and statistically significant in both wealth equations. In the non-financial wealth equation, this coefficient is about half as much as it was in the 2004 sample and in the financial wealth equation it is much larger reflecting the very strong growth in the value of the stock market over 2004 to 2007. However, consumption no longer appears to play a part in the adjustment process as the error correction term does not enter the consumption equation with a statistically significant coefficient. Finally, lagged changes in financial and non-financial wealth predict current consumption growth as in the 2004 sample. The properties of the residuals are somewhat worse for the full sample model as we may expect given the lack of cointegration. There is evidence of both skewness and excess kurtosis in the residuals from both wealth equations and for ARCH effects in the

consumption and financial wealth residuals. In summary, while the direct tests reject cointegration in the full sample, the statistical significance of the error-corrections term in the wealth equations of the VEC model indirectly suggest a cointegrating relationship.

#### *Permanent and Transitory Components*

The moving average representation of the VEC model can be used to decompose the time series of each variable into its permanent component which is the stochastic trend in the series and its transitory component which is the time series of the deviations from its stochastic trend. Specifically, the vector of permanent components is given as

$$X_t^p = X_0 + C(1) \sum_{i=1}^t (\varepsilon_i + \mu) \quad (7)$$

where  $X_0$  is the vector of initial values of the series and  $C(1)$  is the long-run impact matrix of the residuals from the VEC model.<sup>17</sup> Given cointegration, the long-run impact matrix has reduced rank. The vector of transitory components is given as  $X_t^s = X_t - X_t^p$  and is stationary. This decomposition does not require an identification of the underlying structural shocks since it involves the residuals from the estimated VEC model which is a reduced-form model.

The permanent-transitory decomposition in figure 4 is based on the VEC model estimated for the 2004 sub-sample. In the figure the vertical axis marks 2004:3; the end of the estimation period. It is evident from the graphs that non-financial wealth has a larger and more persistent transitory component than any of the other series. For consumption, labor income and financial wealth the estimated transitory components are all relatively small. Thus despite the statistically significant error correction term for the consumption equation (see table 3), its transitory component is of a similar magnitude to

that for labor income and financial wealth. The relatively small transitory component in non-housing consumption suggests that Australian households did not significantly change their consumption in response to transitory fluctuations in housing wealth in the period up to 2004.

The above results are reinforced if the VEC model is used to analyze the out-of-estimation period 2004:4-2008:3. To compute a permanent-transitory decomposition for this period, a time series of residuals to be used in (7) is required for the out-of-estimation period. The residuals are obtained as the difference between the actual value of the series and the one step ahead forecast of the series from the estimated sub-sample model using the actual data. As figure 4 shows the rise in the transitory component of non-financial wealth that begins in 2002-03 continues over the remainder of the sample and is considerably larger than that seen in the earlier period. Non-housing consumption, however, does not respond except perhaps very mildly towards the end of the sample. Noticeably, labor income rises above trend over the out of estimation period, reflecting the strong growth in labor income that occurred over the last four years of the sample. Finally financial wealth remained on trend despite the Australian stock market boom over 2004-2007.

Figure 5 shows the transitory component in each series obtained from the VEC model estimated over the full sample. The most striking feature of the figure is that each series, including consumption, has increased dramatically above trend since 2004. If we take this result at face-value, it points to a marked change in consumption behavior by Australian households. At least two interpretations of the results seem possible. One explanation (following Buiter (2008a,b)) is that the unprecedented doubling of the value

of the stock market over 2004-2007 together with the strong growth in house prices represents a stochastic bubble in asset prices which shows up in the large transitory deviations above trend in both financial and non-financial wealth. Concurrent with the stock market boom, Australia's terms of trade rose dramatically and this may account for the strong increase above trend in labor income seen in the figure. In response to these developments, non-housing consumption rose strongly above trend, contrary to what was seen in earlier periods.

An alternative interpretation of figure 5 is that it provides further evidence of the failure of the Lettau-Ludvigson framework as an adequate model for Australian consumption, income and wealth after 2004. Identifying the exact reason for this failure is difficult. However the apparent failure of cointegration at the end of 2004 points to intrinsic parameter instability in the model for the reasons given by Carroll et al (2006) or the omission of a non-stationary bubble term not accounted for in the derivation of (2) leading to model misspecification. In the latter case it is possible that the previous cointegrating relationship may reassert itself once the bubble has burst.

Whatever the case, we feel the most reliable inferences concerning the relationship between non-housing consumption and non-financial is likely to be found in the data up to 2004. Thus in the remaining empirical analysis we focus on the results for this period.

#### *Dynamic Responses*

In a VEC model, the existence of cointegration allows for identification of the underlying structural shocks as either permanent or transitory. There are several ways to

transform the reduced-form errors ( $\varepsilon_t$ 's) from the VEC model into shocks which have permanent and transitory effects. Following Gonzalo and Ng (2001) define the matrix

$$G = \begin{bmatrix} \alpha'_{\perp} \\ \beta' \end{bmatrix},$$

where  $\alpha_{\perp}$  is the orthogonal complement of  $\alpha$ . It is straightforward to show that the shocks  $\alpha'_{\perp}\varepsilon_t$  (of which there are three here) and  $\beta'\varepsilon_t$  (of which there is only one) have permanent and transitory effects on the levels of the series, respectively. These shocks are in general mutually correlated and are made mutually orthogonal by applying the transformation  $v_t = H^{-1}G\varepsilon_t$  where  $H$  is the lower triangular matrix such that  $HH' = G\Omega G'$ . Thus the orthogonalized permanent and transitory shocks are exactly identified and represent underlying structural shocks. The matrix  $H$  is not unique and it will depend on the ordering of the variables.

The order of the series chosen here is  $X_t = (y_t \ c_t \ a_{ft} \ a_{nt})'$ . Labor income is placed first as the estimated coefficient on the error-correction term in its equation is not significant. The contemporaneous response of the series  $X_t$  to the orthogonal permanent and transitory shocks is given by  $G^{-1}Hv_t$ . Because the matrix  $G$  is in general not lower triangular, the  $i$ th ordered series  $X_{it}$  can respond contemporaneously to the  $j$ th structural innovation  $v_j$ , even if  $j > i$ . So, for example, consumption which is ordered second can potentially respond to all the structural innovations  $v_t$ , not just the first, and this was the case in the reported results.<sup>18</sup>

Figure 6 shows the responses of the series in levels to a one-standard error orthogonal transitory shock in the model estimated to 2004:3 as there is strong evidence

of a VEC representation in this sample. Also shown is the response of the error-correction term. There is a large increase in non-financial wealth in response to the transitory shock and the peak response occurs around three quarters after the initial shock. The response then gradually declines so that by nine quarters non-financial wealth has returned to its initial level. Thereafter it falls somewhat below its initial level prior to complete convergence.<sup>19</sup> By contrast the other series do not respond by nearly as much. Financial wealth increases initially but its response is only about one-third of that of non-financial wealth. Consumption initially falls and then increases somewhat along with labor income. Their peak response is small, occurring at around six quarters, and thereafter both gradually decline together. The error-correction term falls in response to the transitory shock and then gradually increases to zero. Its response primarily reflects the response of non-financial wealth as it is non-financial wealth which adjusts the most to offset deviations from the cointegrating relation.<sup>20</sup>

Table 4 reports the percentage contribution to the forecast-error variance in the growth rate of each series which is attributable to the permanent shocks collectively and to the transitory shock at various forecast horizons. The transitory shock accounts for 38 percent of the forecast-error variance in non-financial wealth at the one-quarter horizon and around 35 percent at longer horizons. This shock is also important for consumption as, along with non-financial wealth, consumption adjusts to deviations from the cointegrating relationship. It accounts for around 35 percent of the forecast-error variance in consumption growth at the one-quarter horizon and for 40 percent at the four-quarter horizon and above. Although the percentage contributions of the transitory shock to consumption and non-financial wealth are similar, it should be noted that consumption

growth is far less volatile than growth in non-financial wealth so that the forecast-error variance is much larger for growth of non-financial wealth than of consumption. The transitory shock is relatively unimportant in explaining fluctuations in both financial wealth and labor income growth. It contributes less than 12 percent to the forecast-error variance in financial wealth growth and under approximately 5 percent in labor income growth at all forecast horizons.

#### *Asset Return Predictability*

The estimated VEC models provide evidence that growth in non-financial wealth responds to lagged  $ca_f a_n y$ . This suggests that  $ca_f a_n y$  should have predictive content for the returns to non-financial wealth and possibly to financial wealth as well. To investigate this implication, the predictive power of  $ca_f a_n y$  for the returns on housing and stocks are evaluated in this section, respectively.<sup>21</sup> The return on housing is calculated from data provided by the Real Estate Institute of Australia.<sup>22</sup> The return on stocks is defined as the quarterly return on the Australian Stock Exchange (ASX) 200 stock market accumulation index. Our results are reported for excess returns defined as quarterly returns on houses and stocks in excess of the 3-month yield to maturity on bank accepted bills.<sup>23</sup>

Table 5 reports the results from the regression of the  $h$ -quarter cumulative excess return on a constant and  $ca_f a_n y$  for housing (Panel A) and for stocks (Panel B) in both the full and sub-samples. The results for the full sample are based on  $ca_f a_n y$  constructed from the DOLS estimates for the sub-sample 1976:4-2004:3; however, the inferences are unchanged if  $ca_f a_n y$  is constructed from the full sample DOLS estimates. Note that results from the full sample regressions may be spurious due to the apparent non-

stationary behavior of  $ca_f a_n y$  in the full sample for either the sub-sample or full sample DOLS estimates. Each panel shows the estimated coefficient on  $ca_f a_n y$  together with its Newey-West  $t$ -statistic and the adjusted R-squared statistic from the regression. Panel A shows that  $ca_f a_n y$  has predictive power for excess returns to housing over the full sample for investment horizons of six to twelve quarters inclusive. For these horizons, the estimated coefficient on  $ca_f a_n y$  is positive and statistically significant and the adjusted R-squared statistic reaches a maximum of 14 percent at eight quarters.<sup>24</sup> The results for the sub-sample are stronger as the  $t$ -statistics are generally more significant and the regressions have more explanatory power. The adjusted R-squared statistic increases from 18 percent at the four quarter horizon to reach around 24 percent at horizons of six to eight quarters and then declines to 13 percent at twelve quarters.

Panel B of the table shows that there is no predictive content in  $ca_f a_n y$  for excess stock returns over the full sample. At no investment horizon is the estimated coefficient on  $ca_f a_n y$  statistically significant and the adjusted R-squared statistics are all negligible. However, there is significant predictive content in  $ca_f a_n y$  for stocks in the sub-sample over investment horizons of one to six quarters inclusive. For these horizons, the estimated coefficient on  $ca_f a_n y$  is positive and statistically significant. The adjusted R-squared statistic increases from 7 percent at the one quarter horizon to 14 percent at four quarters and then declines to 9 percent at six quarters. To summarize, the evidence presented here indicates that over the sub-sample  $ca_f a_n y$  has predictive content for excess returns to housing over medium term investment horizons (one to three years) and for stocks over shorter term horizons (one quarter to around one and half years). There is

also evidence that  $ca_f a_n y$  predicts housing returns in the full sample though this may be spurious. Finally, there is no predictive power in  $ca_f a_n y$  for stock returns over the full sample.

#### **4. Conclusion**

The cointegration approach of Lettau and Ludvigson (2004) is used to investigate the relationship between household non-housing consumption and wealth in Australia. This approach works well in Australian data from 1976:4 to 2004:3. For this sample, we can confidently conclude that household non-financial wealth contains an important and relatively persistent transitory component. At the same time, household consumption was affected by transitory shocks but their impact was much smaller in magnitude than for non-financial wealth. In effect a lot of short to medium term variation in housing wealth never got transmitted to consumption in this sample. This suggests the mechanisms that potentially might link non-housing consumption to changes in housing wealth such as direct wealth effects, the collateral channel or even speculative bubbles did not play an important role in the Australian economy over this time period.

However, in the full sample of data that ends in 2008:3 we cannot reach a definitive conclusion about the responsiveness of non-housing consumption to transitory fluctuations in non-financial wealth. Over 2004-2008, the Australian economy experienced strong growth in its terms of trade and in the value of its stock market. A correction began in mid-to-late 2008, at the end of our sample, in the wake of the global financial crisis. It appears that these developments may have led to the breakdown of the cointegrating relationship for consumption, income and wealth in the Australian economy and to possibly a changed association between non-housing consumption and housing

wealth.

## Appendix

This appendix describes the derivation of the log-linear budget constraint shown by equation (1) in the text. This development draws on Lettau and Ludvigson (2004), Fisher and Voss (2004) and Hamburg, Hoffman and Keller (2005). Aggregate household wealth is assumed to evolve according to the following accumulation equation

$$W_{t+1} = (1 + R_{wt+1})(W_t - C_t). \quad (\text{A1})$$

The variable  $W_t$  is aggregate wealth including human wealth  $H_t$  and non-human wealth  $A_t$ . Private consumption is  $C_t$  and the return on aggregate wealth is  $R_{wt+1}$ . Campbell and Mankiw (1989) derive an expression for the log of the consumption to wealth ratio from a log-linearization of equation (A1). This expression, after dropping linearization constants, is

$$c_t - w_t \approx E_t \sum_{i=1}^{\infty} \rho_w^i (r_{wt+i} - \Delta c_{t+i}), \quad (\text{A2})$$

where  $c_t = \ln(C_t)$ ,  $w_t = \ln(W_t)$ , and  $r_{wt} = \ln(1 + R_{wt})$ . The discount factor is  $\rho_w = 1 - \overline{C_t/W_t}$ , where  $\overline{C_t/W_t}$  is the sample average consumption-wealth ratio, assumed to be constant.

The next step is to decompose the log of total wealth into its components of log human wealth  $h_t$ , log financial wealth  $a_{ft}$  and log non-financial wealth  $a_{nt}$ . It is straightforward to demonstrate the claim of Hamburg et al (2005) that the log of the sum of these components is approximately equal to the sum of the log of these components (apart from a constant term which we drop). That is,

$$w_t \approx \beta_y h_t + \beta_f a_{ft} + \beta_n a_{nt}, \quad (\text{A3})$$

where  $\beta_y = \overline{H_t/W_t}$ ,  $\beta_f = \overline{A_{ft}/W_t}$  and  $\beta_n = \overline{A_{nt}/H_t}$  are the sample averages assumed to be constant. The sample average  $\overline{H_t/W_t}$  is designated as  $\beta_y$  since human capital will be related to labor income. By construction,  $\beta_f + \beta_n + \beta_y = 1$ .

Now derive an expression for the return components. Note that  $R_{wt}$  is implicitly defined as

$$(1 + R_{wt}) = (1 + R_{ht}) \frac{H_t}{W_t} + (1 + R_{aft}) \frac{A_{ft}}{W_t} + (1 + R_{ant}) \frac{A_{nt}}{W_t}, \quad (\text{A4})$$

where  $R_{ht}$  is the return to human wealth and  $R_{aft}$  and  $R_{ant}$  are the return to financial and non-financial wealth, respectively. Following Campbell (1996), it can be shown that after taking the logs of both sides of equation (A4), the approximate relation

$$r_{wt} \approx \beta_y r_{ht} + \beta_f r_{aft} + \beta_n r_{ant} \quad (\text{A5})$$

is obtained, where  $r_{ht} = \ln(1 + R_{ht})$ ,  $r_{aft} = \ln(1 + R_{aft})$  and  $r_{ant} = \ln(1 + R_{ant})$ . Substitute equations (A3) and (A5) into equation (A2) to obtain

$$c_t - \beta_y h_t - \beta_f a_{ft} - \beta_n a_{nt} \approx E_t \sum_{i=1}^{\infty} \rho_w^i (\beta_y r_{ht+i} + \beta_f r_{aft+i} + \beta_n r_{ant+i} - \Delta c_{t+i}). \quad (\text{A6})$$

Finally, define human capital as the present value of current and future after-tax labor income  $Y_{t+j}$  so that

$$H_t = Y_t + E_t \sum_{j=1}^{\infty} \prod_{i=1}^j (1 + R_{ht+i})^{-1} Y_{t+j}. \quad (\text{A7})$$

The accumulation equation corresponding to equation (A7) is

$$H_{t+1} = (1 + R_{ht+1})(H_t - Y_t), \quad (\text{A8})$$

which is analogous to equation (A1). Log-linearize equation (A8) to obtain the expression

$$y_t - h_t \approx E_t \sum_{i=1}^{\infty} \rho_h^i (r_{ht+i} - \Delta y_{t+i}), \quad (\text{A9})$$

where  $y_t = \ln(Y_t)$ ,  $\rho_h = 1 - \overline{Y_t}/\overline{H_t}$  and  $\overline{Y_t}/\overline{H_t}$  is the sample average labor income to human wealth ratio, assumed to be constant. Equation (A9) is analogous to equation (A2). Substitute equation (A9) for  $h_t$  in equation (A6), assume  $\rho_h = \rho_w = \rho$  and recall  $\beta_y = 1 - \beta_f - \beta_n$ , to obtain

$$c_t - \beta_f a_{ft} - \beta_n a_{nt} - \beta_y y_t \approx E_t \sum_{j=1}^{\infty} \rho^j (\beta_y \Delta y_{t+j} + \beta_f r_{aft+j} + \beta_n r_{ant+j} - \Delta c_{t+j}),$$

which is equation (1) in the text.

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**Table 1**

Cointegration tests and long-run estimates

A. Cointegration Tests					
Sample	$PP$	$Z_\tau$	$ADF(4)$	5% C.V.	10% C.V.
1976:4-2008:3	-3.104	-3.094	-2.299	-4.11	-3.83
1976:4-2004:3	-5.017	-4.995	-3.565	"	"
1976:4-1998:4	-4.059	-4.043	-3.165	"	"
B. DOLS Estimates of the cointegrating parameters					
Sample	$\beta_0$	$\beta_f$	$\beta_n$	$\beta_y$	$AR(1)$
1976:4-2008:3	-0.5589 (-16.24)	0.0995 (6.39)	0.1281 (3.81)	0.6970 (9.26)	0.781
1976:4-2004:3	-0.8737 (-17.63)	0.0302 (1.93)	0.2196 (6.36)	0.8161 (13.27)	0.705
1976:4-1998:4	-0.7712 (-8.03)	0.0343 (1.55)	0.2085 (5.09)	0.7696 (9.63)	0.527

*Notes:* This table reports residual based tests for cointegration based on the OLS estimates of the parameters in the cointegrating regressions.  $PP$  and  $Z_\tau$  are the Phillips-Perron (1988) and Phillips-Ouliaris (1990)  $t$ -statistics, respectively. In each case, the autocovariance function is truncated at four lags.  $ADF(4)$  is the adjusted Dickey-Fuller (1981)  $t$ -statistic from the cointegrating regression, including four lags of the first-differences of the OLS residuals. The critical values are taken from Hamilton (1994), table B.9. In the DOLS regressions, the number of leads and lags of changes in the right-hand side variables is  $k=4$  and the standard errors to form the  $t$ -statistics shown in parentheses are based on the Newey-West estimator (Newey and West, 1987) with eight lags. The first-order autocorrelation coefficient of the residuals from the DOLS regressions is denoted  $AR(1)$ .

**Table 2****Contributions to Predicted Consumption Growth**

Sample	Contributions to Predicted Consumption Growth			Consumption Growth	
	Financial	Non-Financial	Labor Income	Predicted	Actual
1976:4-2004:3	0.1	0.7	1.0	1.8	1.7
2002:1-2004:3	0.1	1.8	2.1	4.0	3.7

*Notes:* Contributions to growth are calculated based upon the estimates of the cointegrating parameters presented in panel B of table 1 for the sample 1976:4-2004:3. All numbers are in annual percent.

**Table 3**

Estimates of vector error correction models conditional on DOLS estimates.

Equation	1977:4-2004:3				1977:4 – 2008:3			
	$\Delta c_t$	$\Delta y_t$	$\Delta a_{ft}$	$\Delta a_{nt}$	$\Delta c_t$	$\Delta y_t$	$\Delta a_{ft}$	$\Delta a_{nt}$
Const	0.002 (1.57)	0.001 (0.27)	<b>0.010</b> (3.71)	0.003 (1.13)	0.001 (1.34)	0.001 (0.61)	<b>0.010</b> (3.26)	0.004 (1.66)
$\sum_{i=1}^3 \Delta c_{t-i}$	0.065 (0.34)	<b>0.689</b> (1.99)	-0.361 (-0.79)	-0.574 (-1.20)	-0.030 (-0.16)	<b>0.702</b> (2.02)	-0.603 (-1.14)	-0.542 (-1.13)
$\sum_{i=1}^3 \Delta y_{t-i}$	-0.149 (-0.90)	<b>-0.703</b> (-2.32)	0.632 (1.59)	0.729 (1.73)	0.029 (0.21)	<b>-0.642</b> (-2.59)	0.712 (1.88)	0.250 (0.73)
$\sum_{i=1}^3 \Delta a_{ft-i}$	<b>0.148</b> (2.16)	-0.024 (-0.19)	0.033 (0.20)	0.030 (0.17)	0.108 (1.90)	-0.010 (-0.09)	0.287 (1.80)	0.197 (1.36)
$\sum_{i=1}^3 \Delta a_{nt-i}$	<b>0.154</b> (2.80)	<b>0.296</b> (2.96)	-0.155 (-1.18)	<b>0.605</b> (4.38)	<b>0.207</b> (4.16)	<b>0.279</b> (3.05)	-0.180 (-1.29)	<b>0.479</b> (3.79)
$ec_{t-1}$	<b>-0.162</b> (-2.47) [0.02]	0.007 (0.06) [0.96]	0.153 (0.97) [0.34]	<b>0.504</b> (3.03) [0.00]	-0.014 (-0.32) [0.75]	0.001 (0.01) [0.99]	<b>0.247</b> (2.00) [0.05]	<b>0.251</b> (2.24) [0.03]
$\bar{R}^2$	0.171	0.178	-0.037	0.223	0.145	0.162	0.034	0.216
<i>s.e.</i>	0.008	0.014	0.019	0.020	0.008	0.014	0.021	0.019
<i>sc(4)</i>	0.728	0.436	0.048	0.073	0.518	0.686	0.174	0.165
<i>Reset</i>	0.794	0.886	0.461	0.014	0.475	0.921	0.108	0.121
<i>ARCH</i>	0.100	0.892	0.100	0.536	0.023	0.938	0.016	0.701
<i>Normality</i>	0.923	0.779	0.128	0.000	0.812	0.563	0.000	0.000
<i>Skewnwss</i>	0.703	0.533	0.176	0.348	0.645	0.408	0.032	0.017
<i>Kurtosis</i>	0.918	0.758	0.148	0.000	0.666	0.516	0.000	0.000

*Notes:* In each equation, the sums of the estimated coefficients on the lags of the variables are reported, together with the  $t$ -statistic for the sum.  $t$ -statistics are shown in parentheses. Also shown is the associated  $p$ -value in square brackets for the estimated error correction coefficient ( $ec$ ) which is constructed from the dynamic ordinary least squares (DOLS) estimates reported in table 1. Significant coefficients at the five percent level are highlighted in bold face. The adjusted  $R$ -squared statistic and the standard error of estimate ( $s.e.$ ) are given. Also shown are the  $p$ -values for, respectively: an  $F$ -test for absence of serial correlation of order four ( $sc(4)$ ), based on regression of residuals on initial regressors and four lagged residuals; a reset  $F$ -test for heteroskedasticity based on regression of squared residuals on a constant and fitted values and their squares and cubes; an  $F$ -test for fourth-order ARCH effects based on regression of squared residuals on lagged squared residuals; the Jarque-Bera test for normality and a  $z$ -test for symmetry and absence of excess kurtosis.

**Table 4**

Decomposition of forecast error variance

<i>j</i>	1977:4 – 2004:3							
	$\Delta c_{t+i} - E_t(\Delta c_{t+i})$		$\Delta y_{t+i} - E_t(\Delta y_{t+i})$		$\Delta a_{ft+i} - E_t(\Delta a_{ft+i})$		$\Delta a_{nt+i} - E_t(\Delta a_{nt+i})$	
	<i>P</i>	<i>T</i>	<i>P</i>	<i>T</i>	<i>P</i>	<i>T</i>	<i>P</i>	<i>T</i>
1	65.52	34.48	100.0	0.00	91.31	8.69	62.13	37.87
2	62.16	37.84	99.97	0.03	91.51	8.49	64.87	35.13
3	57.72	42.28	96.64	3.36	91.18	8.82	67.40	32.60
4	60.31	39.69	96.22	3.78	89.61	10.39	66.67	33.33
6	60.31	39.69	95.44	4.56	88.26	11.74	65.50	34.50
8	60.48	39.52	95.15	4.85	88.24	11.76	64.63	35.37
12	60.43	39.57	94.95	5.05	88.16	11.84	64.62	35.38
$\infty$	60.43	39.57	94.94	5.06	88.13	11.87	64.63	35.37

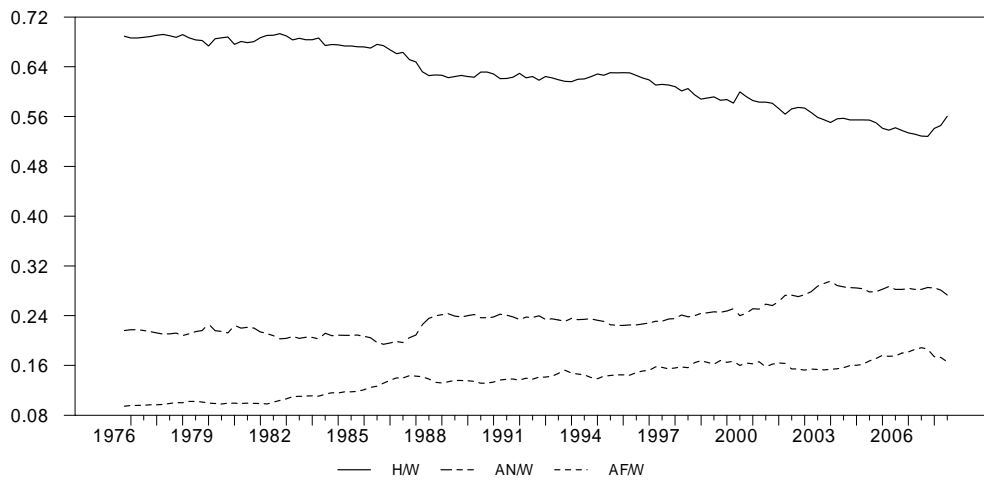
*Notes:* The table reports the percentage contribution of the permanent shocks together (*P*) and the transitory shock (*T*) to the *j*-step ahead forecast error variance in the growth of each series, respectively, from the 2004 VEC model. Horizons are in quarters.

**Table 5** Long-horizon regressions

Regressand $ca_f a_n y_t$	Forecast Horizon (in quarters)									
	1	2	3	4	5	6	8	12	16	20
Panel A. Excess Returns to Housing										
1982:1 – 2008:3	0.173 (1.09) [0.006]	0.401 (1.38) [0.028]	0.587 (1.39) [0.041]	0.963 (1.61) [0.076]	1.286 (1.74) [0.097]	1.728 (2.02) [0.125]	2.371 (2.42) [0.143]	3.197 (3.20) [0.115]	3.294 (1.77) [0.056]	4.152 (1.44) [0.053]
1982:1 – 2004:3	0.321 (1.02) [0.004]	1.044 (1.83) [0.056]	1.578 (1.82) [0.087]	2.736 (2.35) [0.178]	3.436 (2.56) [0.209]	4.205 (3.04) [0.242]	4.826 (3.44) [0.227]	4.750 (3.01) [0.127]	3.375 (1.21) [0.036]	4.610 (1.30) [0.056]
Panel B. Excess Stock Returns										
1977:1 – 2008:3	0.718 (1.34) [0.026]	1.049 (1.37) [0.031]	1.363 (1.46) [0.034]	1.539 (1.39) [0.029]	1.533 (1.19) [0.019]	1.326 (0.94) [0.006]	0.315 (0.19) [-0.008]	-3.310 (-1.31) [0.017]	-4.392 (-1.11) [0.021]	0.148 (0.03) [-0.010]
1977:1 – 2008:3	1.986 (2.35) [0.068]	3.095 (2.50) [0.093]	4.314 (2.66) [0.116]	5.557 (2.63) [0.137]	5.980 (2.47) [0.121]	5.887 (2.29) [0.092]	3.783 (1.36) [0.021]	-2.687 (-0.58) [-0.001]	-2.484 (-0.43) [-0.003]	1.124 (0.16) [-0.010]

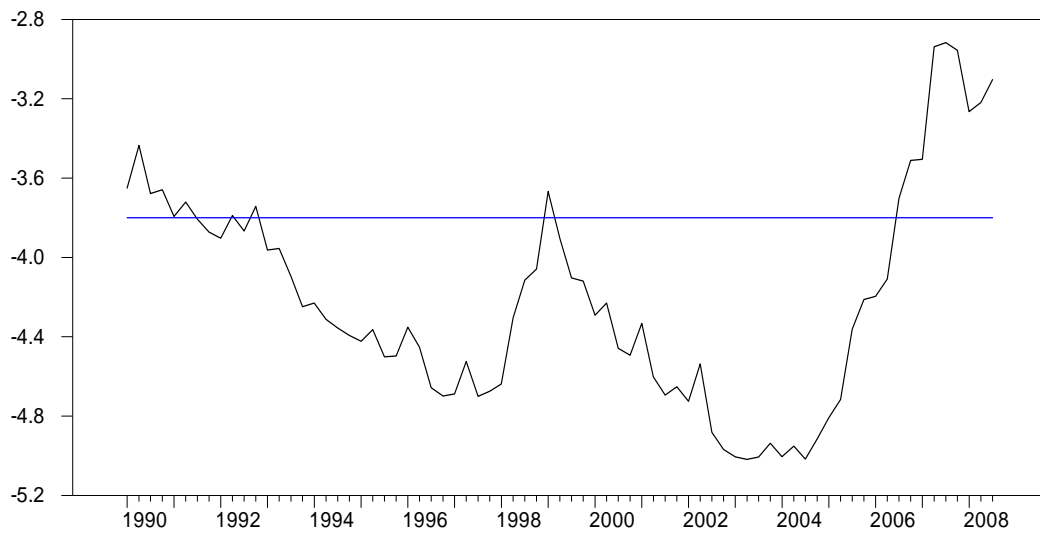
Notes: Panel A reports the results of the OLS regression of the excess return on housing constructed from data provided by the Real Estate Institute of Australia over  $h$  quarters on  $ca_f a_n y_t$ . Panel B reports the results of the OLS regression of the excess return on the ASX200 accumulation index over  $h$  quarters on  $ca_f a_n y_t$ . The housing and stock return over  $h$  quarters is defined as  $\Pi_0^h (1 + r_{t+i})^{-1}$ , where  $r_t$  is the return in excess of the rate on 3-month bank accepted bills. A constant is included in all regressions. For each regression, the estimated coefficient on  $ca_f a_n y_t$  is reported together with the Newey-West  $t$ -statistic (calculated with four lags) directly below in parentheses. The adjusted R-squared statistic from the regression is shown in square brackets. In every case, the coefficients in  $ca_f a_n y$  are the DOLS estimates for the sample that ends in 2004:3 shown in table 1.

**Figure 1. Wealth Shares**



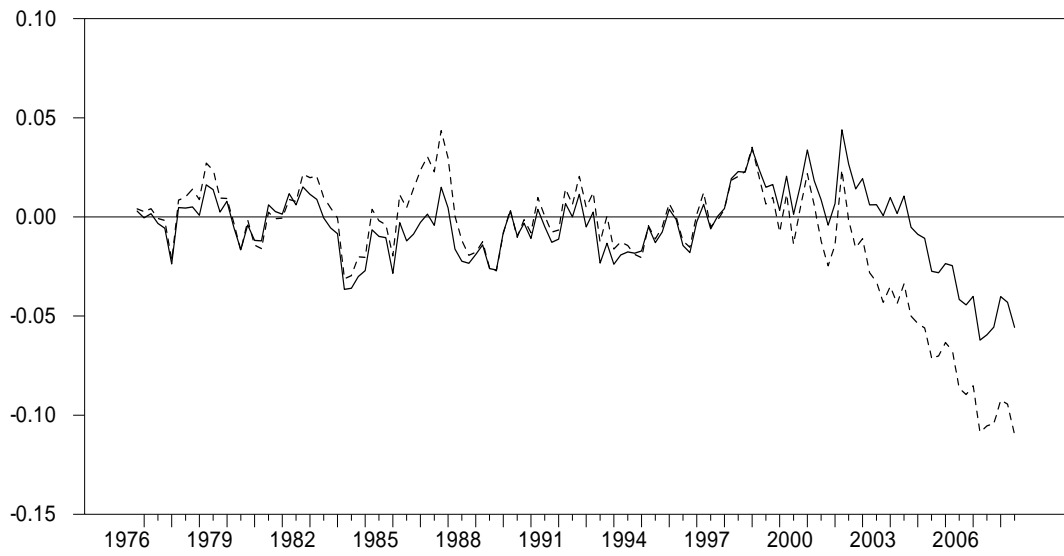
*Notes:* Total wealth ( $W$ ) is defined as the sum of human wealth ( $H$ ), non-financial wealth ( $AN$ ) and financial wealth ( $AF$ ). The proxy for human wealth is labor income discounted by a quarterly return of 2.5 percent. The figure shows each component of wealth as a proportion of total wealth.

**Figure 2.** Phillips-Perron test for cointegration in recursive samples



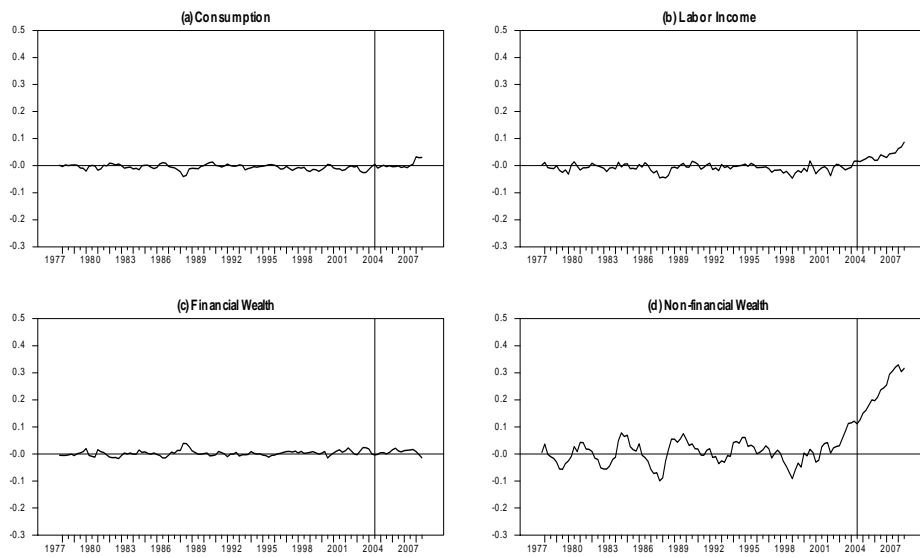
*Notes:* The figure shows the Phillips-Perron (1988) test statistic in recursive samples. The 10 percent critical value of  $-3.84$  is shown by the horizontal line.

**Figure 3.** DOLS cointegrating residuals



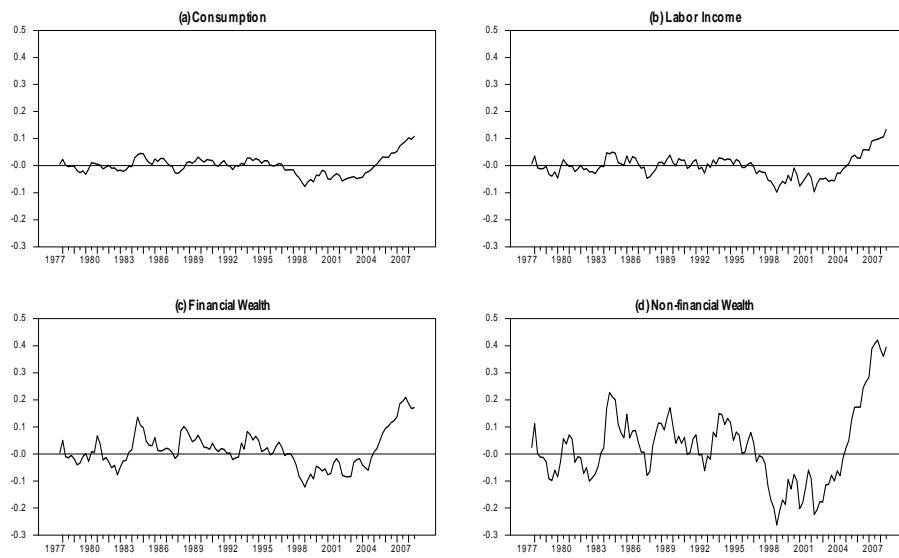
*Notes:* The figure shows the DOLS cointegrating residuals for the full sample using the full sample cointegrating relation (shown as the solid line) and the sub-sample cointegrating relation to 2004:3 (shown as the dashed line).

**Figure 4.** Transitory component of each series in sub-sample model



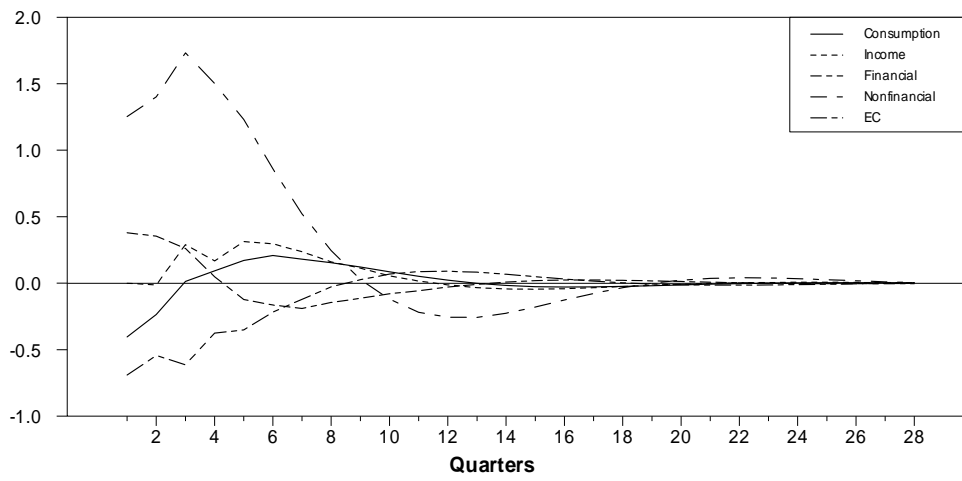
*Notes:* The transitory components are from the VEC model estimated to 2004:3 under the restriction that labor income is weakly exogenous. This end-of-estimation date is shown by the vertical line.

**Figure 5.** Transitory component of each series in full sample model



*Notes:* The transitory components are from the VEC model estimated over the full sample under the restriction that labor income is weakly exogenous.

**Figure 6.** Responses of series to a transitory shock in sub-sample model.



*Notes:* The figure shows the responses to a one-standard error transitory shock. The VEC model is estimated to 2004:3 under the restriction that labor income is weakly exogenous.

## Footnotes

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<sup>1</sup> Other studies that have addressed this issue following the approach of Lettau and Ludvigson include; Fernandez-Corugedo, Price and Blake (2003) for United Kingdom, Pichette and Tremblay (2003) for Canada, Hamburg, Hoffman and Keller (2005) for Germany, De Veirman and Dunstan (2008) for New Zealand and Chen (2006) for Sweden.

<sup>2</sup> Lettau and Ludvigson's finding is broadly consistent with the predictions of the permanent income hypothesis. Increases or decreases in future wealth that are expected to be transitory have only a modest effect on permanent income and hence on consumption.

<sup>3</sup> Evidence that U.S. household asset wealth is dominated by equity market returns is reflected in the very high correlation between quarterly fluctuations in household net worth and movements in equity returns (Lettau and Ludvigson, 2004).

<sup>4</sup> Australia experienced rapid increases in house prices in the late 1980s and again in 2002 and 2003.

<sup>5</sup> The importance of the wealth effect on consumption in Australia has previously been investigated by Tan and Voss (2003). However, their analysis is conducted in a single-equation framework which, as they note, assumes all of the adjustment to changes in wealth occurs through consumption changes and is likely to over-estimate the wealth effect on consumption.

<sup>6</sup> Strictly speaking with time-varying real asset returns the approximate Euler equation for consumption growth is  $E_t \Delta c_{t+1} = \mu + \theta E_t r_{j,t+1}$  for  $j = af$  or  $an$ ; however log consumption will be close to a random walk with drift for small  $\theta$  i.e. when the inter-temporal elasticity of substitution is small.

<sup>7</sup> A summary of calculating returns to human capital and suitable values is provided in Rosen (2008).

<sup>8</sup> On the basis of the PP and ADF unit root tests, the null of a unit root in each series cannot be rejected at standard significance levels so there may exist a cointegrating relationship among them.

<sup>9</sup> This finding could be due to an omitted variable that has become important recently. One possibility is the relative price of housing which has increased recently. It is defined as the implicit price deflator of household expenditure on rent and other dwelling services divided by the implicit price deflator of total household consumption expenditure excluding expenditure on rent and other dwelling services. Inclusion of this variable in the cointegrating regressions had no effect on the magnitude of the test statistics and cannot account for the breakdown of cointegration in the full sample.

<sup>10</sup> Hansen uses the fully modified cointegration estimator to perform the *SupF* test while we employ DOLS.

<sup>11</sup> Davidson and MacKinnon (1993) argue that these single equation methods have low power to reject the null hypothesis of no cointegration. If this is the case then actually rejecting the null on the basis of these tests would seem to provide very strong evidence of cointegration.

<sup>12</sup> We also used the systems approach of Johansen (1991) to test for the number of cointegrating relationships. For the full sample, both the trace and  $\lambda$  - max statistics fail to reject the null of no cointegration at the 10 percent level. However, the trace and  $\lambda$  -

max statistics reject the null of no cointegration at the 10 percent level for the sample to 1998:4 and reject only marginally above the 10 percent level for the sample to 2004:3. There was no evidence for a second cointegrating vector in either sample.

<sup>13</sup> The DOLS standard errors are corrected for heteroskedasticity and serial correlation using the Newey and West (1987) estimator.

<sup>14</sup> The ratio of the log of total consumption to the log of total consumption excluding expenditure on rent and other dwelling services increased steadily from 1976:4 to 1993:3 and thereafter declined slowly mirroring the behaviour of rent and other dwelling services as a proportion of total household consumption. This brings into question the assumption of a constant proportional relationship.

<sup>15</sup> For both sub-samples, the Johansen method gave an estimate of the coefficient on non-financial wealth that was somewhat larger than the DOLS estimate. The estimate of the coefficient on financial wealth was similar to the DOLS estimate in the sample to 2004:3 but was very small and negative in the sample to 1998:4. For the full sample, the estimated coefficient on financial wealth was again a small negative value and the coefficient on non-financial wealth was considerably larger than the DOLS estimate.

<sup>16</sup> The DOLS coefficient estimates imply long-run MPC's out of wealth (at the mid-point 1992:3) of 0.0101 for financial wealth and 0.0076 for non-financial wealth in the full sample. For the 2004 sub-sample, the implied MPC out of financial wealth is considerably smaller (0.0031) while the implied MPC out of non-financial wealth is considerably larger (0.0130).

<sup>17</sup> Formally,  $C(1) = I + \sum_{i=1}^{\infty} C_i L^i = \beta_{\perp} \gamma \alpha'_{\perp}$  where  $L$  is the lag operator,  $\beta_{\perp}$  and  $\alpha_{\perp}$  are the orthogonal complements of  $\alpha$  and  $\beta$  respectively,  $\gamma = (\alpha'_{\perp} \Psi \beta_{\perp})^{-1}$  and  $\Psi = I - \sum_{i=1}^{k-1} \Pi_i$ .

Here the vector of permanent components corresponds to the permanent components in the multivariate Beveridge-Nelson decomposition under cointegration. As a practical matter, Gonzalo and Ng (2001) recommend constraining statistically insignificant error-correction coefficients in estimated  $\alpha$  to zero before constructing  $\alpha_{\perp}$  because the long-run impact matrix of the  $\varepsilon_t$ 's (which depends on  $\alpha_{\perp}$ ) can be very sensitive to small variations in estimated  $\alpha$ . In this case, the variable is said to be weakly exogenous with respect to the parameters in the cointegrating relationship. In all the results that follow, we constrained the coefficient on the error-correction term in the labor income equation to zero because labor income was the only variable which appeared consistently weakly exogenous. However, the results were unaffected by imposing this restriction.

<sup>18</sup> Alternatively, the contemporaneous interactions among the series in the orthogonal VEC model (re-parameterized to a model in levels of the series) is given by  $H^{-1}GX_t$ . Consumption potentially depends on contemporaneous labor income, financial wealth and non-financial wealth because the matrix  $G$  is in general not lower triangular.

<sup>19</sup> For the ordering of the variables chosen here, the source of the transitory shock can be attributable to fluctuations in non-financial wealth only if the other variables in the VEC model are weakly exogenous, that is, only if the error correction term does not appear in the equations for growth in labor income, consumption and financial wealth. To see this, let  $C(1)$  and  $\Gamma(1)$  denote the long-run impact matrix of the reduced form shocks ( $\varepsilon_t$ 's) and

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the orthogonal shocks ( $v_t$ 's), respectively. The relationship between them is  $\Gamma(1) = C(1)G^{-1}H$ . Using the results in Fisher and Huh (2007), one can establish that the last column of  $C(1)$  will be a column vector of zeros and  $G^{-1}H$  will be lower triangular provided growth in labor income, consumption and financial wealth are weakly exogenous. It then follows that the last column of  $\Gamma(1)$  is a column vector of zeros so that the orthogonal innovation in non-financial wealth has no long-run impact on any of the variables; it only has a transitory effect. It is not possible to identify the transitory shock exclusively with fluctuations in non-financial wealth here as not all the other variables are weakly exogenous in the VEC model.

<sup>20</sup> Confidence intervals around the impulse responses are not shown because figure 6 would appear too cluttered. One-standard error bands were calculated by taking the estimated coefficients in the VEC model (both for the short-run dynamics and the long-run relation) to form the data generating process which was then bootstrapped 1000 times. The response of non-financial wealth to the transitory shock is statistically significant up to horizons of eight quarters. The response of financial wealth becomes statistically insignificant after two quarters. Both the consumption and labor income responses are statistically significant from around five to eight quarters and the negative consumption responses seen initially are also statistically significant. The impulse responses with one-standard error bands are available on request.

<sup>21</sup> The predictive content of  $cay$  (not  $ca_f a_n y$ ) for stock returns was first investigated by Lettau and Ludvigson (2001) for the United States and subsequently by other researchers for other countries. Among those studies are; Fernandez-Corugedo, Price and Blake (2003) for the United Kingdom, Fisher and Voss (2004) for Australia and Ioannidis et al. (2006) for Australia, Canada and the United Kingdom.

<sup>22</sup> The Real Estate Institute of Australia provides data for the quarterly median house price and the quarterly median rent of three bedroom homes in each major Australian city. From this data, we calculated the median quarterly return (both capital and rental return) on houses in each major city. We form an Australia-wide return by taking a weighted average of the city returns. The weights are calculated as the ratio of the state population, which is concentrated in the major city, to the Australia-wide population.

<sup>23</sup> The quarterly excess return on stocks of *minus* 55% in 1987:4, which corresponds to the stock market crash in October of that year, is a highly influential observation. This observation is set to zero through a regression of excess returns on a dummy variable which takes the value of one in 1987:4 and zero otherwise. The results for real returns are very similar to those for excess returns and are available on request.

<sup>24</sup> In the long horizon regressions, the estimated coefficient on  $ca_f a_n y$  is expected to be positive. If the representative household expects the returns on housing and stocks to increase in the future, it will temporarily increase consumption above the long run trend it shares with labor income and the two wealth components. Thus,  $ca_f a_n y$  rises in anticipation of an increase in returns implying a positive coefficient.