Consumption, (Dis)Aggregate Wealth and Asset Returns

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Abstract

In this work, I analyze the importance of the disaggregation of asset wealth into its main components (financial and housing wealth). I show, from the consumer's intertemporal budget constraint, that the residuals of the trend relationship among consumption, financial wealth, housing wealth and labor income (summarized by the variable *cday*) should help to predict quarterly asset returns, and to provide better forecasts than a variable like *cay* from Lettau and Ludvigson (2001), which considers aggregate wealth instead.

Using data for the United Kingdom, I show that the superior forecasting power of *cday* is due to: (*i*) its ability to track the changes in the composition of asset wealth and the specificities of the different assets; and (*ii*) the faster rate of convergence of the coefficients to the "long-run equilibrium" parameters.

Unlike Lettau and Ludvigson (2001, 2004), the results suggest that, while financial wealth shocks are mainly transitory, fluctuations in housing wealth are very important due to their persistence. Governments and central banks should, therefore, pay special attention to the behavior of housing markets (and to a smaller extent to the behavior of financial markets) when defining macroeconomic stabilizing policies.

Keywords: financial wealth, housing wealth, consumption, expected returns.

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1 Introduction

A growing body of empirical literature has documented the long-term predictability of asset returns.[1]

One important reason for the interest in the linkages between wealth and other macroeconomic variables is that expected excess returns on assets appear to vary with the business cycle (Lettau and Ludvigson, 2001). Different explanations have been offered, namely: inefficiencies of financial markets; the rational response of agents to time-varying investment opportunities driven by cyclical variation in risk aversion (Sundaresan, 1989; Constantinides, 1990; Campbell and Cochrane, 1999) or in the joint distribution of consumption and asset returns.

Lettau and Ludvigson (2001) introduced a new approach to investigate these linkages and have shown that the transitory deviation from the common trend in consumption, aggregate wealth and labor income, *cay*, is a strong predictor of asset returns, as long as the expected return to human capital and consumption growth are not too volatile. Fernandez-Corugedo *et al.* (2003) use the same approach, but incorporate the relative price of durable goods, showing that unless the relative price of durables and non-durables is constant, it needs to be taken into account in modelling. More recently, Julliard (2004) shows that the expected changes in labor income (which capture the movements in human capital) also carry relevant information for predicting future asset returns, because of their ability to track time varying risk premia.

In this paper, I use the representative consumer intertemporal budget constraint to derive an equilibrium relation between the transitory deviation from the common trend in consumption, housing wealth, financial wealth and labor income and expected future asset returns, and show that the consumption-(dis)aggregate wealth ratio, *cday* provides better forecasts than *cay*.

A simple formulation of the life cycle model suggests that consumers spread the increases in anticipated wealth over time and that the wealth effects on consumption should be the same in magnitude whichever is the component of wealth considered.

However, the responsiveness of consumers to financial and housing wealth shocks can be different for several reasons:[2] (*i*) differences in liquidity; (*ii*) utility derived from the property right of an asset,

[1] See, for example, Fama and French (1988), Campbell and Shiller (1988), Poterba and Summers (1995), Richards (1995), Lettau and Ludvigson (2001, 2004).

[2] For a more detailed discussion, see Case *et al.* (2001). Note, however, that the empirical evidence in this area is still inconclusive. Elliott (1980), Levin (1998) and Mehra (2001) found that the wealth effect is independent of the category of wealth. Thaler (1990), Sheiner (1995), and Hoynes and McFadden (1997) investigated the correlation between individual saving rates and changes in house prices and found a weak relation. In contrast, Case (1992), Kent and Lowe (1998), Skinner (1999), Case *et al.* (2001), and Dvornak and Kohler (2003) found evidence of a considerable housing wealth effect

such as housing services or bequest motives; (*iii*) different distributions of assets across income groups; (*iv*) expected permanency of changes of different categories of assets; (*v*) mismeasurement of wealth;[3] and (*vi*) 'psychological factors'.

First, housing assets and financial assets have different degrees of liquidity: if agents can purchase and sell assets with different liquidities, then the chosen consumption-wealth ratio is not independent of the timing of income payments (Pissarides, 1978). For instance, housing is often considered a 'lumpy' asset, because it may be difficult to liquidate only a part of it, and transaction costs tend to be high. This implies that the coefficient on housing wealth should be lower than that on stock market wealth.[4]

Second, housing represents both an asset and a consumption item. When house prices increase, wealth may increase, but so does the cost of housing services (Poterba, 2000). This factor makes it less likely that increased wealth in housing is consumed, resulting in a lower marginal propensity to consume out of housing wealth.[5] On the other hand, households may have different motives about bequeathing their stock portfolios and bequeathing their homesteads to heirs or may view the accumulation of some kinds of wealth as an end in itself: for instance, house ownership provides a visible sign of status.

Third, while housing wealth tends to be held by consumers in all income classes, financial wealth, on the other hand, is in many countries concentrated in the high-income groups which are often thought to have a lower propensity to consume. In this case, changes in housing wealth might have a larger impact on consumption than changes in financial wealth.

Fourth, consumers may view increases in wealth for some asset groups as more likely to be permanent, while others are more likely to be viewed as temporary or uncertain. This difference in perception of the permanency of price changes can be related to past experiences of sudden price reversals in asset markets and implies that if an increase (or decrease) in wealth is seen as permanent, it is more likely to increase (or decrease) long-run consumption.

Fifth, consumers may not be able to accurately measure wealth, especially for houses which are less homogenous and less frequently traded than shares. Many consumers may also not be aware of the exact value of their indirect share holdings, such as pension funds, until they are close to retirement

on consumption.

[3] This may be especially so for houses which are less homogenous and less frequently traded than shares. Also many consumers may not be aware of the exact value of their indirect share holdings. For example, Sousa (2003) shows that directly held stock market wealth effects are significantly different from indirectly held stock market wealth effects.

[4] Note, however, that financial innovations, such as the availability of home equity loans, are likely to increase the liquidity of housing assets (Muellbauer and Lattimore, 1999).

[5] A related argument is based on the idea that for every household that sells a house there is a household that buys it. Therefore, in aggregate, the increase in the seller's consumption could be offset by the decrease in buyer's consumption (Bajari *et al.*, 2003).

age. For example, Sousa (2003) shows that directly held stock market wealth effects on consumption are significantly different from indirectly held stock market wealth effects.

Finally, consumers may attach certain psychological factors to specific assets. Shefrin and Thaler (1988) show that consumers may use 'mental accounts' and earmark certain assets as more appropriate to use for current expenditure, while others are reserved for long-term savings.

Each of these explanations suggests a distinction between the impact of housing wealth and financial wealth on consumption. Therefore, I argue that the disaggregation of wealth is an important issue and should also be considered in the context of forecasting future asset returns. This follows from the fact that the consumption-(dis)aggregate wealth ratio summarizes agent's expectations of future returns on assets and consumption growth: when average asset returns (this is, housing asset returns and financial asset returns) are expected to be higher (lower) in the future, forward-looking investors will increase (decrease) consumption, allowing it to rise (decrease) above (below) its common trend with housing wealth, financial wealth, and labor income. In this way, investors may insulate future consumption from fluctuations in expected returns. This is particularly important since the composition of wealth is very different across countries and governments and central banks frequently take into account the behavior of both types of assets when defining macroeconomic policies.[6]

The rest of the paper is organized as follows. In Section 2, I present the theoretical framework linking consumption, financial wealth, housing wealth, labor income and expected returns and how I express the important predictive components of the consumption-(dis)aggregate wealth ratio in terms of observable variables. In Section 3, I briefly present the data, estimate the model and discuss the results. Using data for the United Kingdom, I show that financial wealth effects are significantly different from housing wealth effects. Unlike Lettau and Ludvigson (2001, 2004), who argue that asset wealth fluctuations are largely transitory (and, therefore, not important for economic policy considerations), the results suggest that, while substancial fluctuations in financial assets need not indeed be associated with large subsequent movements in consumption, fluctuations in housing assets are very important due to their persistence. An important implication is that governments and central banks need to pay special attention to the behavior of housing markets (and to a smaller extent to the behavior of financial markets) when defining macroeconomic stabilizing policies. In Section 4, I test the implication that deviation from trend relationship among consumption, (dis)aggregate wealth and labor income, *cday*, are likely to lead asset returns. I show that the superior forecasting power of *cday* is due to: (*i*) its ability to track the changes in the composition of asset wealth and the specificities of the different assets; and

[6] See, for example, Banks *et al.* (2002) for a comparison of wealth porfolios in the U.K. and in the U.S. and Bertaut (2002) for a discussion about the evolution of the composition of wealth across countries.

(ii) the faster rate of convergence of the coefficients to the "long-run equilibrium" parameters. Finally, in Section 5, I conclude and refer the main limitations of the model and the lines of direction for future research.

2 The Consumption-(Dis)Aggregate Wealth ratio

Consider a representative agent economy in which all wealth, including human capital, is tradable. Let W_t be aggregate wealth (human capital plus asset holdings) in period t. C_t is consumption and $R_{w,t+1}$ is the net return on aggregate wealth. The equation for the accumulation of aggregate wealth may be written:[7]

$$W_{t+1} = (1 + R_{w,t+1})(W_t - C_t) \tag{1}$$

Define $r := \log(1 + R)$, and use lowercase letters to denote log variables throughout. Campbell and Mankiw (1989) show that, if consumption-aggregate wealth ratio is stationary, the budget constraint may be approximated by taking a first-order Taylor expansion of (1). The resulting approximation gives an expression for the log difference in aggregate wealth

$$\Delta w_{t+1} \approx k + r_{w,t+1} + (1 - 1/\rho_w)(c_t - w_t) \tag{2}$$

where ρ_w is the steady-state ratio of new investment to total wealth, $(W - C)/W$, and k is a constant that plays no role in the analysis.[8] Solving this difference equation forward and imposing that $\lim_{i\to\infty} \rho_w^i \left(c_{t+i} - w_{t+i}\right) = 0$, the log consumption-wealth ratio may be written as

$$c_t - w_t = \sum_{i=1}^{\infty} \rho_w^i (r_{w,t+i} - \Delta c_{t+i}). \tag{3}$$

Equation (3) holds not only ex-post (as a consequence of agent's intertemporal budget constraint), but also ex-ante. Accordingly, take conditional expectations of both sides of (3) to obtain

$$c_t - w_t = E_t \sum_{i=1}^{\infty} \rho_w^i (r_{w,t+i} - \Delta c_{t+i}), \tag{4}$$

where E_t is the expectation operator conditional on information available at time t. Equation (4) shows that, if the consumption-aggregate wealth ratio is not constant, it must forecast changes in asset returns or in consumption growth, this is, it can only vary if consumption growth or returns or both are predictable.

[7] Labor income does not appear explicitly in this equation because of the assumption that the market value of tradable human capital is included in aggregate wealth.

[8] We omit unimportant linearization constants in the equations from now on.

Because aggregate wealth (in particular, human capital) is not observable, this framework is not directly suited for predicting asset returns. To overcome this obstacle, Lettau and Ludvigson (2001) assume that the nonstationary component to human capital, denoted H_t, can be well described by aggregate labor income, Y_t, implying that $h_t = k + y_t + z_t$, where k is a constant and z_t is a mean zero stationary random variable.[9]

(Dis)Aggregate wealth can be decomposed as

$$W_t = F_t + U_t + H_t, \tag{5}$$

where F_t is financial asset holdings and U_t is housing asset holdings. This last last expression can be approximated as an expression for the log (dis)aggregate wealth

$$w_t \approx \alpha_f f_t + \alpha_u u_t + (1 - \alpha_f - \alpha_u) h_t, \tag{6}$$

where α_f and α_u equal, respectively, the share of financial asset holdings in total wealth, F/W, and the share of housing asset holdings in total wealth, U/W.

The return to (dis)aggregate wealth can be decomposed into the returns of its components

$$1 + R_{w,t} \approx \alpha_f (1 + R_{f,t}) + \alpha_u (1 + R_{u,t}) + (1 - \alpha_f - \alpha_u)(1 + R_{h,t}). \tag{7}$$

Campbell (1996) shows that (7) maybe transformed into an approximation equation for log returns taking the form

$$r_{w,t} \approx \alpha_f r_{f,t} + \alpha_u r_{u,t} + (1 - \alpha_f - \alpha_u) r_{h,t}. \tag{8}$$

Substituting (6) and (8) into the ex-ante budget constraint (4) gives

$$c_t - \alpha_f f_t - \alpha_u u_t - (1 - \alpha_f - \alpha_u) h_t = E_t \sum_{i=1}^{\infty} \rho_w^i \left\{ [\alpha_f r_{f,t+i} + \alpha_u r_{u,t+i} + (1 - \alpha_f - \alpha_u) r_{h,t+i}] - \Delta c_{t+i} \right\}. \tag{9}$$

This equation still contains the unobservable variable h_t on the left-hand side. To remove it, the formulation linking the log of labor income to human capital, $h_t = k + y_t + z_t$, is replaced into (9),

[9] This assumption may be rationalized by a number of different specifications. First, labor income may be described as the annuity value of human wealth, $Y_t = R_{h,t+1} H_t$, where $R_{h,t+1}$ is the net return of human capital. In this case, $r_{h,t} \equiv \log(1 + R_{h,t+1}) \approx 1/\rho_y (y_t - h_t)$, where $\rho_y \equiv (1 + Y/H)/(Y/H)$, implying $z_t = -\rho_y r_{h,t}$. Second, one could specify a "Gordon growth model" for human capital by assuming that expected returns to human capital are constant and labor income follows a random walk, in which case z_t is a constant equal to $\log(R_h)$. Finally, aggregate labor income can be thought of as the dividend on human capital, as in Campbell (1996) and Jagannathan and Wang (1996). In this case, the return to human capital may be fixed as $1 + R_{h,t+1} = (H_{t+1} + Y_{t+1})/H_t$, and a log-linear approximation of $R_{h,t+1}$ implies that $z_t = E_t \sum_{j=0}^{\infty} \rho_h^j (\Delta y_{t+1+j} - r_{h,t+1+j})$. In each of these cases, the log of aggregate labor income captures the nonstationarity component of human capital.

which yields an approximate equation describing the log consumption-(dis)aggregate wealth ratio using observable variables on the left-hand side

$$c_t - \alpha_f f_t - \alpha_u u_t - (1-\alpha_f-\alpha_u)y_t = E_t \sum_{i=1}^{\infty} \rho_w^i \left\{ [\alpha_f r_{f,t+i} + \alpha_u r_{u,t+i} + (1-\alpha_f-\alpha_u) r_{h,t+i}] - \Delta c_{t+i} \right\} + \eta_t, \tag{10}$$

where $\eta_t = (1-\alpha_f-\alpha_u)z_t$.

Since all the terms on the right-hand side of (10) are presumed to be stationary, c, f, u and y must be cointegrated, and the left-hand side of (10) gives the deviation in the common trend of c_t, f_t, u_t, and y_t. The trend deviation term $c_t - \alpha_f f_t - \alpha_u u_t - (1-\alpha_f-\alpha_u)y_t$ is denoted as $cday_t$.[10] Equation (10) shows that $cday_t$ will be a good proxy for market expectations of future financial, $r_{f,t+i}$, and housing asset returns, $r_{u,t+i}$, as long as expected future returns on human capital, $r_{h,t+i}$, and consumption growth Δc_{t+i}, are not too variable, or as long as these variables are highly correlated with expected returns on assets. When the left hand side of equation (10) is high, consumers expect either high future financial asset returns, or high housing asset returns on market wealth or low future consumption growth. Since this equation takes into account the composition of asset wealth, it should provide a better proxy for market expectations of future returns $(r_{f,t+i}, r_{u,t+i})$ and future consumption growth as long as human capital returns are not too variable.

After this presentation, I briefly describe the data, estimate the trend relationship among consumption, financial wealth, housing wealth and labor income, and present the main results, which is done in the next Section.

3 Estimating the Trend Relationship Among Consumption, (Dis)Aggregate Wealth and Income

The methodology adopted for the estimation of the model consists of two stages. First, I estimate the long-run relation among consumption, financial wealth, housing wealth and income. Then, I proceed with the analysis of short-run dynamics using a Vector-Error Correction Model (VECM).

[10] Lettau and Ludvigson (2001) do not consider the issue of wealth disaggregation. Their specification is given by

$$cay_t = E_t \sum_{i=1}^{\infty} \rho_w^i \left\{ [\alpha r_{a,t+i} + (1-\alpha) r_{h,t+i}] - \Delta c_{t+i} \right\} + (1-\alpha) z_t,$$

where cay_t denotes the trend deviation term $c_t - \alpha a_t - (1-\alpha)y_t$, c_t is consumption, a_t is total asset holdings, y_t is labor income, and α is the share of total asset holdings in total wealth.

3.1 Data

In the estimations, I use quarterly, seasonally adjusted data for the United Kingdom and all variables are measured at 2001 prices, and expressed in the logarithmic form of per capita terms. The definition of consumption, excludes durable and semi-durable goods consumption. Data on income includes only labor income. Original data on wealth correspond to the end-period values. Therefore, I lag once the data, so that the observation of wealth in t corresponds to the value at the beginning of the period $t+1$. The main data source is the Office for National Statistics (ONS), although for housing wealth, I also use data from Halifax plc, the Nationwide Building Society and the Office of the Deputy Prime Minister. In Appendix A, I present a detailed discussion of data.

3.2 The long-run relation

I first use the Phillips-Perron (PP) tests[11] to determine the existence of unit roots in the series and conclude that all the series are first-order integrated, $I(1)$. Next, I analyze the existence of cointegration among the series using the methodology of Engle and Granger (1987), and find evidence that supports this hypothesis. The results of the PP tests and the cointegration tests are presented in Appendix B.[12] Finally, I estimate the trend relationship among consumption, wealth and labor income following Davidson and Hendry (1981), Blinder and Deaton (1985), Ludvigson and Steindel (1999), and Davis and Palumbo (2001) among others. However, since the impact of different assets' categories on consumption can be different (Zeldes, 1989; and Poterba and Samwick, 1995), I disaggregate wealth into its main components: financial wealth and housing wealth. Following Saikkonen (1991) and Stock and Watson (1993), I use a dynamic least squares (DOLS) technique, specifying the following equation

$$c_t = \mu + \beta_f f_t + \beta_u u_t + \beta_y y_t + \sum_{i=-k}^{k} b_{f,i} \Delta f_{t-i} + \sum_{i=-k}^{k} b_{u,i} \Delta u_{t-i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t-i} + \varepsilon_t, \tag{11}$$

where the parameters β_f, β_u,β_y represent, respectively, the long-run elasticities of consumption with respect to financial wealth, housing wealth, and labor income and Δ denotes the first difference operator.[13]

[11] The ADF (Augmented Dickey-Fuller tests) generate the same results, although they have lower power.

[12] These methodologies have limitations and Harris (1995) and Maddala and Kim (1998) present a detailed description of the panoply of alternative tests for cointegration.

[13] The parameters β_f, β_u,β_y should in principle equal $R_h F/(Y+R_h F+R_h U)$, $R_h U/(Y+R_h F+R_h U)$ and $Y/(Y+R_h F+R_h U)$, respectively, but, in practice, may sum to a number less than one, because only a fraction of total consumption expenditure is observable (Lettau and Ludvigson, 2001). Because of this, we decided to write β_f, β_u and β_y instead of α_f, α_u and α_y to distinguish long-run elasticities of *our definition* of consumption from long-run elasticities of *total* consumption.

Implementing the regression in (11) using data for the United Kindom in the period 1977:Q4 - 2001:Q1,[14] generates the following estimates (ignoring coefficient estimates on the first differences) for the shared trend among consumption, financial wealth, housing wealth and income:

$$c_t = \underset{(3.31)}{1.37} + \underset{(6.43)}{0.17}\, f_t + \underset{(2.88)}{0.04}\, u_t + \underset{(5.72)}{0.52}\, y_t. \tag{12}$$

where the Newey-West (1987) t-corrected statistics appear below the coefficient estimates.[15]

The estimations show that the long-run elasticity of consumption with respect to financial wealth (0.17) is more than four times greater than the long-run elasticity with respect to housing wealth (0.04), reflecting the importance of this component of wealth and, simultaneously, the significance of the disaggregation of wealth. As expected, the coefficients of equation (12) do not sum to unity, since I exclude from the definition of consumption the durable and semi-durable goods' consumption. However, the average share of this measure of consumption in total consumption in the sample is 76%, which is approximately equal to the sum of the coefficients of equation (12), namely, 73%. Finally, the implied shares, calculated by scaling the coefficients on financial wealth, housing wealth and income by the inverse sum of the coefficients are, respectively, 0.23, 0.06 and 0.71, which are very plausible figures, since they correspond, approximately, to shares of capital and labor of 0.29 and 0.71, respectively.

3.3 The short-term dynamics

I proceed with the analysis of how consumption reacts to shocks on wealth and how this deviation from the long-run relation is corrected. I want to determine whether deviations from the shared trend in consumption, financial wealth, housing wealth and income are better described as transitory movements in financial wealth and/or housing wealth or as transitory movements in consumption and labor income.

The estimated model is specified as follows:

$$\Delta \mathbf{X}_t = \theta + \gamma_t \beta_t' \mathbf{X}_{t-1} + \Gamma(L)\Delta \mathbf{X}_{t-1} + e_t, \tag{13}$$

[14] As an additional issue of the estimation, I analyze the stability of the cointegrating vector using the methodology of Seo (1998) and splitting the sample in subsamples. The results suggest that the cointegrating vector is relatively stable over time and if there is a structural break, this is close to the beginning point of the sample, at the time of the oil shocks. This is in contrast with Lettau and Ludvigson (2004), who argue that for the U.S. the sample instability comes from the large appreciations of the stock markets during the nineties.

[15] We experimented with various lead/lag lengths in estimating the DOLS specification. For the results reported in (12), we use the value of $k = 1$. However, neither the cointegrating parameter estimates nor the forecasting results we present below are sensitive to the particular value of k. In the case of the consumption-wealth ratio, cay_t, it is computed as $cay_t = c_t - 0.12a_t - 0.83y_t$, where c_t is consumption , a_t is total asset holdings, and y_t is labor income. For the U.S., Lettau and Ludvigson (2001) compute cay_t as $c_t - 0.31a_t - 0.59y_t$.

where $\mathbf{X}_t = (c_t, f_t, u_t, y_t)$ is the vector of consumption, financial wealth, housing wealth, and labor income, $\gamma_t = (\gamma_c, \gamma_f, \gamma_u, \gamma_y)'$ is a $(4x1)$ vector, $\beta_t = (1, -\beta_f, -\beta_u, -\beta_y)'$ is the vector of estimated cointegration coefficients shown in equation (12), and $\Gamma(L)$ is a finite-order distributed lag operator. Thus, γ_t is the short-run adjustment vector telling us how the variables react to the last period's cointegrating error while returning to long-term equilibrium after a deviation occurs; β_t measures the long-run elasticities of one variable respective to another; the term $\beta_t' \mathbf{X}_{t-1}$ measures the cointegrating residual, $cday_{t-1}$. Table 1 presents the results of the estimation using a one-lag cointegrated VAR.[16]

Table 1: Estimates from a Cointegrated VAR.

	Equation			
Dependent variable	Δc_t	Δf_t	Δu_t	Δy_t
Δc_{t-1}	-0.211***	0.359	0.408**	-0.379**
(t-stat)	(-1.870)	(0.487)	(2.082)	(-2.430)
Δf_{t-1}	-0.003	-0.052	-0.010	0.020
(t-stat)	(-0.195)	(-0.467)	(-0.332)	(0.828)
Δu_{t-1}	0.041	0.020	0.777*	0.177*
(t-stat)	(1.186)	(0.087)	(12.790)	(3.665)
Δy_{t-1}	0.135***	0.672	0.237***	-0.014
(t-stat)	(1.876)	(1.422)	(1.890)	(-0.143)
θ	0.123	-1.991*	0.007	-0.413*
(t-stat)	(1.282)	(-3.170)	(0.041)	(-3.115)
$\widehat{cday}_{t-1}$	-0.086	1.467*	-0.005	0.307*
(t-stat)	(-1.222)	(3.192)	(-0.045)	(3.162)
$\bar{R}^2$	0.065	0.103	0.709	0.139

This table reports the estimated coefficients from cointegrated vector-autoregressions (VAR).

Symbols *, **, *** represent, respectively, significance level of 1%, 5% and 10%.

Newey-West (1987) corrected t-statistics appear in parenthesis. The sample period is 1977:Q4 to 2001:Q1.

The table reveals some interesting properties of the data on consumption, financial wealth, housing wealth, and labor income.[17] First, estimation of the consumption growth equation shows that

[16] The lag length was chosen in accordance with findings from Akaike and Schwarz tests.

[17] As an additional issue of the estimation, I also analyze the stability of the short-term adjustment vector and the presence of an asymmetric behavior in the response of consumption to different wealth shocks. The results suggest that

$\widehat{cday}_{t-1}$ does not predict consumption growth. The sign of the coefficient is negative and its value (approximately, -0.09) is small, suggesting that the correction is very slow. On the other hand, consumption growth is somewhat predictable by the lag of consumption growth as noted by Flavin (1981), Campbell and Mankiw (1989), which can be interpreted as a sign of some delay in the adjustment of consumption. The lagged values of labor income growth are also statistically significant, which may follow from habit persistence, near-rational rules of thumb, or liquidity constraints.[18] Second, estimation of the financial wealth growth equation shows that $\widehat{cday}_{t-1}$ is statistically significant. Moreover, the estimated coefficient (1.467) suggests that $\widehat{cday}_{t-1}$ strongly predicts financial wealth growth and implies that deviations in financial wealth from its shared trend with consumption, housing wealth, and labor income uncover a very important transitory variation in financial wealth. Third, estimation of housing wealth growth equation shows that $\widehat{cday}_{t-1}$ does not help to predict housing wealth growth: the estimated coefficient is very small (-0.005) and it is not statistically significant. However, it is shown that the lagged values of consumption growth, of housing wealth growth and of labor income growth are statistically significant. Moreover, the $\bar{R}^2$ statistic shows that this equation explains more than 70% of the housing wealth growth.

In sum, these results suggest that deviations from the shared trend in consumption, financial wealth, housing wealth, and labor income are mainly described as transitory movements in financial wealth. In the other hand, changes in house wealth contain an important persistent component and are not responsible for most of the short-term adjustment. Therefore, when consumption deviates from its habitual ratio with financial wealth, housing wealth and labor income, it is financial wealth that is forecast to adjust until the equilibrating relationship is restored; forward-looking households foresee changes in the return of their future financial wealth. This is in contrast with Lettau and Ludvigson (2001, 2004) who argue that total asset wealth changes are mainly transitory. In fact, the results suggest than only the financial component of asset wealth change is transitory.

the short-term adjustment vector remains relatively stable over time and that there is no evidence of an asymmetric behaviour.

[18] This evidence differs from the results of Lettau and Ludvigson (2001), who find that only lagged consumption growth is significant.

4 Does the (Dis)Aggregation of Wealth help to predict better Asset Returns and Consumption Growth?

I have argued that significant loading of the long-run relationship among consumption, (dis)aggregate wealth and income reflects agents' expectations of future changes in asset returns or consumption growth - in accordance with equation (10). Moreover, since I disaggregate asset wealth into its main components (financial and housing wealth) and take, therefore, into account the different composition and specificities of the asset holdings, I argue that $\widehat{cday}_t$ should provide a better forecast than a variable like $\widehat{cay}_t$ in Lettau and Ludvigson (2001).

4.1 Forecasting quarterly asset returns

I look at total asset returns - namely, the MSCI - UK Total Return Index - for which quarterly data are available and should provide a good proxy for nonhuman components of asset wealth. I denote r_t the log real return of the index in consideration and $r_{f,t}$ the log real yield rate of 3-month Treasury Bill (the "risk-free" rate). The log excess return is $r_t - r_{f,t}$.

Figures 1 and 2 plot, respectively, the standardized trend deviations, $\widehat{cday}_t$ and $\widehat{cay}_t$, and the excess return on the MSCI - UK Total Return Index over the period spanning 1977:Q4 and 2001:Q1. They show a large diversity of episodes for which $\widehat{cday}_t$ is able to forecast better future asset returns than $\widehat{cday}_t$, namely: the housing market boom of 1977-1979; the stock market crash of 1987; the housing market boom of 1986-1989; most of the stock market fluctuations of the nineties.

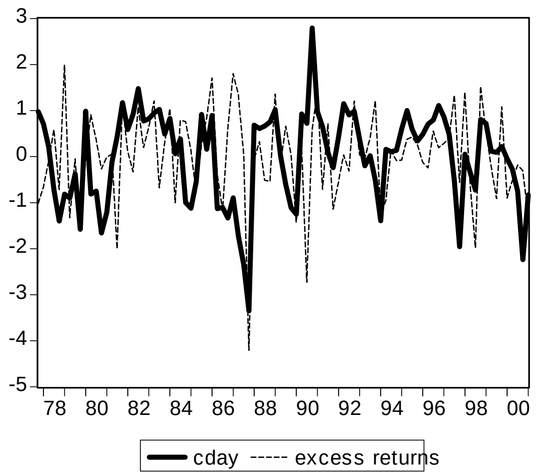


Figure 1: Times series of cday and excess returns.

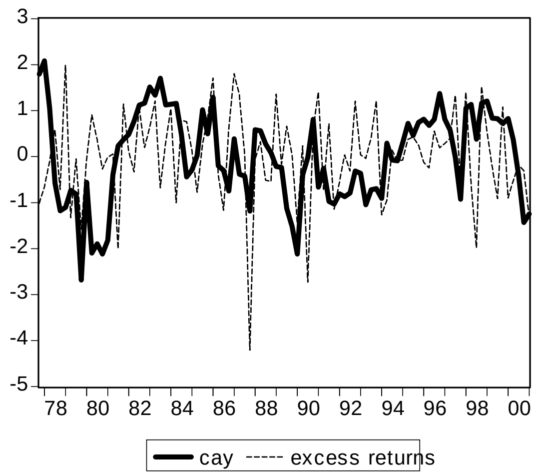


Figure 2: Times series of cay and excess returns.

I now move on to assess the forecasting power of $\hat{cday}_t$ - the deviations of consumption from its trend relationship with financial wealth, housing wealth and income - and to compare it with $\hat{cay}_t$- the deviations of consumption from its trend relationship with aggregate wealth and income -, which is summarized in Table 2. The table reports estimates from OLS regressions of log one-period ahead real returns (Panel A) and excess returns (Panel B) on the variables named at the head of a column.

Table 2: Forecasting quarterly excess returns using $\hat{\text{cday}}$.and $\hat{\text{cay}}$.

Constant (t-stat)	lag (t-stat)	$\hat{cday}_t$ (t-stat)	$\hat{cay}_t$ (t-stat)	$\bar{R}^2$
Panel A: Real Returns				
0.025*	-0.071			0.00
(3.193)	(-1.001)			
-2.153**		1.595**		0.04
(-2.487)		(2.520)		
0.225**			0.893***	0.01
(2.170)			(1.934)	
-2.634**	-0.162***	1.950*		0.06
(-2.468)	(-1.770)	(2.498)		
0.252**	-0.104		1.005***	0.01
(2.190)	(-1.444)		(1.964)	
Panel B: Excess Returns				
0.017**	-0.112			0.00
(2.148)	(-1.460)			
-2.039**		1.505**		0.04
(-2.305)		(2.328)		
0.198**			0.810***	0.01
(2.003)			(1.832)	
-2.638**	-0.202**	1.947**		0.07
(-2.456)	(-2.306)	(2.479)		
0.233**	-0.142***		0.960***	0.02
(2.120)	(-1.863)		(1.947)	

Symbols *, **, *** represent, respectively, significance level of 1%, 5% and 10%.

Newey-West (1987) corrected t-statistics appear in parenthesis.

The sample period is 1977:Q4 to 2001:Q1.

Table 2 shows that the regressions of returns on one lag of the dependent variable (Panel A, for real returns; and Panel B, for excess returns) are quite weak. This model has no forecasting power for both real returns and excess returns.

By contrast, the trend deviation explains an important fraction of the variation in next quarter's return. It is shown that $\widehat{cday}$ helps to predict better future returns than $\widehat{cay}$: in both the estimation of excess returns and real returns, $\widehat{cday}$ explains 4% of the variation in next quarter, while $\widehat{cay}$ explains only 1%. The predictive impact of $\widehat{cday}$ on future returns is economically larger than that of $\widehat{cay}$: the point estimate of the coefficient on $\widehat{cday}$ is about 1.595 for real returns (0.893 in the case of $\widehat{cay}$) and about 1.505 for excess returns (0.810 in the case of $\widehat{cay}$). Thus, a one-standard-deviation increase in $\widehat{cday}$ leads to, approximately, a 132.92 basis points rise in the expected real return on MSCI - UK Total Return Index and a 125 42 basis points increase in the expected excess returns, this is, respectively, a 5.42% and a 5.11% increase at an annual rate. On the other hand, $\widehat{cay}$ itself has a standard deviation of about 0.015, implying that a one-standard-deviation increase in $\widehat{cay}$ leads to, approximately, a 59.5 basis points rise in the expected real return on MSCI-UK Total Return Index and a 54 basis points increase in the expected excess returns, this is, respectively, a 2.40% and a 2.18% increase at an annual rate.

Finally, regressions of real returns and excess returns on their own lags and on one lag of trend deviation, produce roughly the same results as the previous regressions.

These results accord well with the economic intuition from the framework presented in Section 2. If returns on assets are expected to decline in the future, investors who desire to smooth consumption paths will allow consumption to fall temporarily below its long-term relationship with financial wealth, housing wealth and labor income in an attempt to insulate future consumption from lower returns, and vice versa. Thus, investors' optimizing behavior suggests that deviations in the long-term trend among c, f, u and y should be positively related to future asset returns.

4.2 Long-horizon forecasts

I also examine the relative predictive power of $\widehat{cday}$ for returns at longer horizons and compare it with $\widehat{cay}$. In principle, $\widehat{cday}$ could be a long-horizon forecaster of consumption growth, asset returns, or both.[19] Tables 7, 8, and 9 present the results of single-equation regressions of consumption growth, and real returns and excess returns, over horizons spanning 1 to 4 quarters, on trend deviation $\widehat{cday}$ and compare them with $\widehat{cay}$. In the estimation of the regressions of consumption growth, the dependent

[19] For a discussion on the empirical proxies for the consumption-wealth ratio and their forecasting power see Hahn and Lee (2005) and Rudd and Whelan (2006)

variable is the H-period consumption growth rate $\Delta c_{t+1} + ... + \Delta c_{t+H}$; in the estimation of the regressions of excess returns, the dependent variable is the H-period log excess return on the MSCI - UK Total Return Index, $r_{t+1} - r_{f,t+1} + .. + r_{t+H} - r_{f,t+H}$; in the estimation of the regressions of real returns, the dependent variable is the H-period log real return on the MSCI - UK Total Return Index, $r_{t+1} + r_{t+H}$. For each regression, the tables report the estimates from OLS regressions on $\widehat{cday}$ (Panel A) and $\widehat{cay}$ (Panel B).

Consistent with the estimation of the cointegrated VAR summarized in Table 1 and with Lettau and Ludvigson (2001), the results shown in Table 3 suggest that $\widehat{cday}$ has no predictive power for future consumption growth. The individual coefficients are not statistically significant, are small in magnitude and the $\bar{R}^2$ are all close to zero.

Table 3: Long-run horizon regressions for consumption growth.

	Forecast Horizon H			
Regressor	1	2	3	4
Panel A: Consumption Growth, using $\widehat{cday}_t$				
$\widehat{cday}_t$	-0.03	-0.17***	-0.12	-0.17
(t-stat)	(-0.36)	(-1.68)	(-0.67)	(-0.83)
$\bar{R}^2$	[0.00]	[0.03]	[0.00]	[0.01]
Panel B: Consumption Growth, using $\widehat{cay}_t$				
$\widehat{cay}_t$	0.09	0.08	0.21	0.25***
(t-stat)	(1.26)	(0.90)	(1.72)	(1.83)
$\bar{R}^2$	[0.02]	[0.00]	[0.05]	[0.05]

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

Newey-West (1987) corrected t-statistics appear in parenthesis.

The sample period is 1977:Q4 to 2001:Q1.

Table 4 reports results from forecasting of the log real returns on the MSCI - UK Total Return Index. Panel A shows that $\widehat{cday}$ has a significant forecasting power for future real returns, particularly at 3 and 4 quarters horizons, with the $\bar{R}^2$ statistic reaching 0.20. In comparison, Panel B shows that $\widehat{cay}$ performs worse: the coefficient estimates are less statistically significant, smaller in magnitude and, for the same horizons, the $\bar{R}^2$ statistic ranges between 0.12 and 0.14.

Table 5 reports results from forecasting of the log excess returns on the MSCI - UK Total Return Index, which roughly replicate those found in the previous Table.

Table 4: Long-run horizon regressions for real returns.

Regressor	Forecast Horizon H 1	2	3	4
	Panel A: Real Returns, using $\widehat{cday}_t$			
$\widehat{cday}_t$	1.59**	3.13*	4.91*	5.40*
(t-statistic)	(2.52)	(3.31)	(4.11)	(4.17)
$\bar{R}^2$	[0.04]	[0.10]	[0.20]	[0.20]
	Panel B: Real Returns, using $\widehat{cay}_t$			
$\widehat{cay}_t$	0.89***	2.07**	3.21*	3.81*
(t-statistic]	(1.93)	(2.50)	(3.22)	(3.81)
$\bar{R}^2$	[0.01]	[0.06]	[0.12]	[0.14]

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

Newey-West (1987) corrected t-statistics appear in parenthesis.

The sample period is 1977:Q4 to 2001:Q1.

Table 5: Long-run horizon regressions for excess returns.

Regressor	Forecast Horizon H 1	2	3	4
	Panel A: Excess Returns, using $\widehat{cday}_t$			
$\widehat{cday}_t$	1.51**	2.89*	4.61*	5.08*
(t-stat)	(2.33)	(3.02)	(3.88)	(3.85)
$\bar{R}^2$	[0.04)	[0.10]	[0.19]	[0.19]
	Panel B: Excess Returns, using $\widehat{cay}_t$			
$\widehat{cay}_t$	0.81***	1.84**	2.94*	3.56*
(t-stat)	(1.83)	(2.25)	(2.93)	(3.54)
$\bar{R}^2$	[0.01]	[0.05]	[0.10]	[0.13]

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

Newey-West (1987) corrected t-statistics appear in parenthesis.

The sample period is 1977:Q4 to 2001:Q1.

In sum, the results suggest that the disaggregation of wealth into its main components is an important issue in the context of forecasting future asset returns. Not only $\widehat{cday}$ performs better than $\widehat{cay}$, but its relative predictive power is also greater for larger periods. As in Lettau and Ludvigson

(2001), the results also suggest that $\widehat{cday}$ has no predictive power for future consumption growth and that lagged returns do not forecast next quarter's variation both of real returns and excess returns.

4.3 Out-of-sample forecasts

This section compares the forecasting ability of $\widehat{cday}_t$ and $\widehat{cay}_t$ in an out-of-sample context.[20] This exercise faces several econometric issues.

First, Ferson *et al.* (2002) argue, with a simulation exercise, that if both expected returns and the predictive variable are highly persistent the in-sample regression results may be spurious, and both $\bar{R}^2$ and statistical significance of the regressor are biased upward.[21] The autocorrelation of realized returns is low in the data,[22] nevertheless the degree of persistence of expected returns is not observable.[23] On the other hand, since $\widehat{cday}$ and $\widehat{cay}$ are autocorrelated, this could give rise to spurious regression results.[24] As a consequence, in addition to in-sample predictions presented in the previous Section, I also performe out-of-sample forecasts.[25]

Second, a "look-ahead" bias might arise from the fact that the coefficients used to generate $\widehat{cday}_t$ and $\widehat{cay}_t$ are estimated using the full data sample, this is, using a fixed cointegrating vector.[26] To address this issue we also look at out-of-sample forecasts where $\widehat{cday}_t$ and $\widehat{cay}_t$ are reestimated every period, using only the data available at the time of the forecast, and the predictive regressions are estimated recursively using data from the beginning of the sample to the quarter immediately preceding the forecast period. The difficulty with this technique, as argued in Lettau and Ludvigson (2002), is that it can strongly understate the predictive ability of the regressor, which would make it more difficult for $\widehat{cday}$ (and $\widehat{cay}$) to display forecasting power if the theory is true.

With these caveats in mind, I nevertheless compare the forecasting ability of $\widehat{cday}_t$ and $\widehat{cay}_t$ using the Root Mean Squared Error, the Theil's U, the McCracken (2000) MSE-F statistic, and the Clark and McCracken (2001) ENC-NEW statistic.[27]

[20] Foster *et al.* (1997) and Rapach and Wohar (2005) provide a theoretical analysis of data mining in predictive regression models.

[21] See also Torous *et al.* (2005).

[22] The autocorrelation of the realized MSCI-UK returns is -0.11.

[23] The return may be considered to be sum of an unobservable expected return plus a unpredictable noise, and the predictable component could be highly autocorrelated.

[24] This is a common problem for both $\widehat{cday}_t$ and $\widehat{cay}_t$, but is likely to be less severe for the former than the latter since their first autocorrelations are, respectively, 0.55 and 0.71.

[25] Inoue and Kilian (2004) show that in-sample and out-of-sample tests of predictability are, under the null of no predictability, asymptotically equally reliable.

[26] For a discussion on the potential "look-ahead" bias, see Brennan and Xia (2005).

[27] The Theil's U is the ratio of the root-mean-squared errors for the unrestricted and restricted regression model forecasts.

Tables 6 and 7 compare the out-of-sample forecasting power of $\widehat{cday}_t$ (Panel A) and $\widehat{cay}_t$ (Panel B) for real and excess returns over horizons of 1, 2 and 4 quarters, using a fixed cointegrating vector; Tables 8 and 9 repeat the same exercise, but the cointegrating vector is instead reestimated every period using only the data available at the time of the forecast. Moreover, since Brennan and Xia (2002) show that changing the starting point of the out-of-sample forecast might dramatically change the measured performance, I use three different starting points for the out-of-sample forecast: 1987:Q4, 1992:Q4 and 1997:Q4, corresponding, respectively, to the first ten, fifteen and twenty years of available data.

The results shown in Tables 6 and 7 show that $\widehat{cday}_t$ performs better than $\widehat{cay}_t$ in forecasting real and excess returns. It can be seen $\widehat{cday}_t$ has a significant out-of-sample forecasting ability, corroborated by the different statistics used. Moreover, the predictive power also increases substancially as we increase the horizon over which future returns should be predicted, in accordance to the in-sample forecasting power reported in the previous sub-Section.

Tables 8 and 9 provide results which are not so striking, showing that the performance of $\widehat{cday}_t$ is similar to $\widehat{cay}_t$ in forecasting real and excess returns. This is, however, not very surprising since consistent estimation of the parameters requires a large number of observations, and an out-of-sample procedure is likely to induce significant sampling error in the coefficient estimates during the earling estimation recursions, as argued by Lettau and Ludvigson (2002).

If the mean-squared error (MSE) for the unrestricted model forecasts is less than the MSE for the restricted model forecasts, then U <1. In the estimations, the restricted or benchmark model is the model of constant returns.

The MSE-F statistic is a variant of the Diebold and Mariano (1995) and West (1996) statistic and is used to test whether the unrestricted regression model forecasts are significantly superior to the restricted model forecasts.

The ENC-NEW statistic is a variant of the Harvey *et al.* (1998) statistic designed to test for forecast encompassing.

Table 6: Out-of-sample Forecasts of Real Returns: Fixed Cointegrating Vector.

	Forecast Horizon H											
	1				2				4			
	Panel A: Real Returns, using $\widehat{cday}_t$											
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0705	0.656	0.689***	2.431**	0.0915	0.551	2.213***	5.329*	0.1139	0.428	4.950**	9.597*
1992:Q4	0.0653	0.661	1.332***	1.602**	0.0895	0.571	0.166	1.652***	0.1048	0.417	6.873**	6.524**
1997:Q4	0.0859	0.811	-0.842	-0.144	0.1157	0.776	-1.811	-0.528	0.1351	0.686	-2.683	-0.600
	Panel B: Real Returns, using $\widehat{cay}_t$											
First forecast period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0714	0.713	0.591	1.055	0.0926	0.600	1.339	2.747***	0.1192	0.480	4.383***	7.044*
1992:Q4	0.0676	0.686	-1.175	-0.077	0.0934	0.589	-2.876	-0.180	0.1191	0.463	-1.807	2.221
1997:Q4	0.0869	0.772	-1.276	-0.513	0.1195	0.723	-2.751	-1.063	0.1540	0.670	-4.936	-1.681

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

Table 7: Out-of-sample Forecasts of Excess Returns: Fixed Cointegrating Vector.

	Forecast Horizon H											
	1				2				4			
Panel A: Excess Returns, using $\widehat{cday}_t$												
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0717	0.725	0.487	2.051**	0.0915	0.630	1.841***	4.364**	0.1166	0.523	3.881**	7.490**
1992:Q4	0.0650	0.726	1.407***	1.515**	0.0879	0.646	0.410	1.557***	0.1048	0.503	7.102**	6.209*
1997:Q4	0.0858	0.852	-0.641	-0.079	0.1133	0.821	-1.642	-0.491	0.1329	0.747	-2.508	-0.560
Panel B: Excess Returns, using $\widehat{cay}_t$												
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0723	0.791	0.659***	0.909	0.0919	0.690	1.589***	2.334***	0.1199	0.5865	5.448**	6.539**
1992:Q4	0.0672	0.763	-0.929	-0.048	0.0916	0.675	-2.436	-0.192	0.1188	0.5614	-1.209	2.144***
1997:Q4	0.0869	0.832	-1.089	-0.443	0.1169	0.787	-2.512	-0.998	0.1512	0.748	-4.821	-1.668

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

Table 8: Out-of-sample Forecasts of Real Returns: Cointegrating Vector Reestimated.

	Forecast Horizon H											
	1				2				4			
Panel A: Real Returns, using $\widehat{cday}_t$												
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0722	0.739	-14.875	-3.752	0.0964	0.658	-25.297	-2.052	01282	0.549	-38.542	-1.542
1992:Q4	0.0670	0.763	-2.402	-0.895	0.0927	0.711	-4.766	-1.259	0.1192	0.597	-8.614	-0.496
1997:Q4	0.0822	0.835	0.419	0.252	0.1086	0.820	1.093	0.671	0.1153	0.698	3.636***	2.217***
Panel B: Real Returns, using $\widehat{cay}_t$												
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0722	0.742	-16.152	-4.702	0.0963	0.663	-25.270	-6.051	0.1278	0.554	-34.591	-1.248
1992:Q4	0.6706	0.745	-4.775	-1.557	0.0932	0.688	-11.179	-2.776	0.1219	0.579	-17.776	0.166
1997:Q4	0.0832	0.847	-1.431	-0.030	0.1047	0.714	-2.191	1.506	0.0793	0.397	-4.187	7.458***

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

Table 9: Out-of-sample Forecasts of Excess Returns: Cointegrating Vector Reestimated.

	Forecast Horizon H											
	1				2				4			
Panel A: Excess Returns, using $\widehat{cday}_t$												
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0732	0.830	-12.891	-3.650	0.0959	0.775	-19.647	-3.418	0.1300	0.698	-28.997	-4.975
1992:Q4	0.0667	0.867	-2.925	-0.949	0.0913	0.847	-6.374	-1.440	0.1199	0.777	-11.947	-0.580
1997:Q4	0.0823	0.920	0.553	0.353	0.1054	0.916	1.422***	0.934	0.1090	0.802	5.137**	3.400***
Panel B: Excess Returns, using $\widehat{cay}_t$												
First Forecast Period	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW	RMSE	Theil's U	MSE-F	ENC-NEW
1987:Q4	0.0732	0.832	-16.016	-4.786	0.0958	0.780	-23.960	-5.605	0.1298	0.705	-30.387	-0.948
1992:Q4	0.0669	0.837	-5.573	-1.558	0.0917	0.808	-13.740	-2.894	0.1222	0.736	-21.382	0.433
1997:Q4	0.0834	0.833	-1.431	-0.030	0.1023	0.671	-2.191	1.506	0.0781	0.356	-4.187	7.458***

Symbols *, ** and *** represent significance at a 1%, 5% and 10% level, respectively.

In addition to these out-of-sample forecasts, I also perform a very simple exercise using rolling-samples: the coefficients of $\widehat{cday}$ and $\widehat{cay}$ are first estimated using the smallest number of observations; then, one observation is added at each time and the coefficents are recursively estimated. This exercise provides an idea about the rate of convergence of the coefficients to the "long-run equilibrium" coefficients. Figures 3, 4 and 5 plot the pattern of the coefficients of $\widehat{cday}$, while Figures 6 and 7 plot the pattern of the coefficients of $\widehat{cay}$. Despite the instability associated to early estimations, it is clear that the coefficents of $\widehat{cday}$ converge to the "long-run equilibrium" coefficients at a faster rate than $\widehat{cay}$. This is, therefore, an important element that helps to explain the superior forecasting power of $\widehat{cday}$ relative to $\widehat{cay}$.

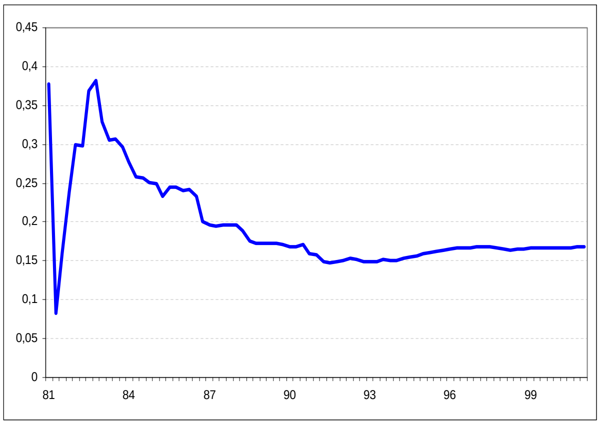


Figure 3: Coefficient associated to financial wealth.

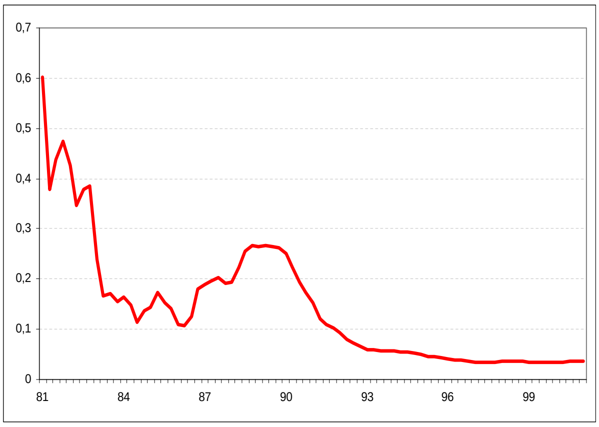


Figure 4: Coefficient associated to housing wealth.

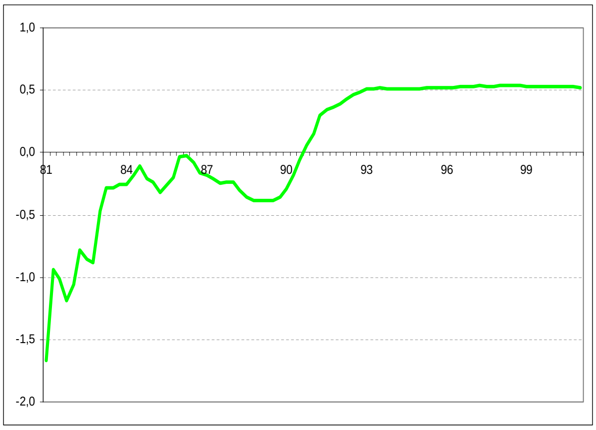


Figure 5: Coefficient associated to labor income.

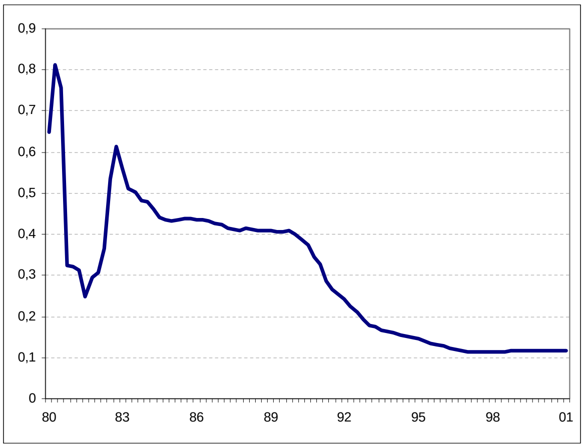


Figure 6: Coefficient associated to asset wealth.

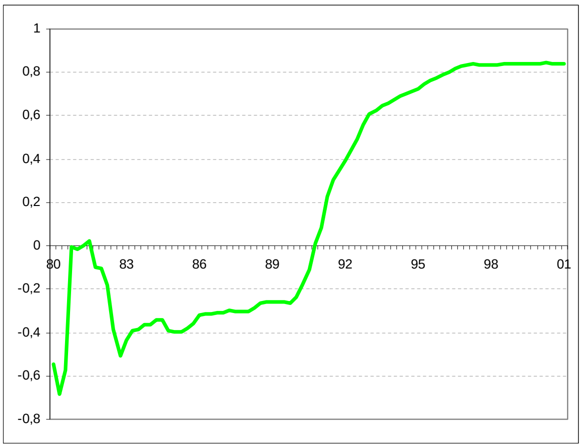


Figure 7: Coefficient associated to labor income.

In sum, the results suggest that $\widehat{cday}$ performs better than $\widehat{cay}$ in an out-of-sample context and that the predictive ability is stronger over longer horizons, corroborating the results found in the in-sample exercise shown in the previous sub-Section. Moreover, it is shown that the rate of convergence of the coefficients of $\widehat{cday}$ to the "long-run equilibrium" coefficients is faster than $\widehat{cay}$. Therefore, the disaggregation of wealth into its main components is an important issue in the context of forecasting future asset returns.

4.4 A skeptical look at the data using a VAR approach

As a robustness check of the previous results, this sub-Section does not impose the theoretical restrictions implied by the budget constraint in equation (10).[28] The results are consistent and show that the joint estimation of the forecasting equations for real returns,[29] consumption growth, financial wealth growth, housing wealth growth, and labor income growth imply that: (*i*) lagged returns do not have forecasting power, but *cday* is an important proxy for the expectations about future asset returns; (*ii*) financial wealth changes are mainly transitory; (*iii*) housing wealth changes are very persistent...

I estimate the following Vector Autoregressive Model (V AR)

$$\mathbf{X}_t = \theta + A(L)\mathbf{X}_{t-1} + \xi_t, \tag{14}$$

[28] In a recent paper, Koop *et al.* (2005) question the key findings of Lettau and Ludvigson (2001, 2004), namely, that most changes in wealth are transitory and have no effect on consumption. The authors use a Bayesian model averaging and argue that there is model uncertainty with regards to the number of cointegrating vectors, the form of deterministic components, lag length and whether the cointegrating residuals affect consumption and income directly.

[29] The same results are obtained when I use instead excess returns.

where $\mathbf{X}_t = (r_t, \Delta c_t, \Delta f_t, \Delta u_t, \Delta y_t)$ is the vector of real returns, consumption growth, financial wealth growth, housing wealth growth, and labor income growth, $\Gamma(L)$ is a finite-order distributed lag operator, and ξ_t is a vector of error terms.[30] For comparison, I also estimate the following VAR that adds $\widehat{cday}_{t-1}$ as an exogenous variable

$$\mathbf{X}_t = \theta + A(L)\mathbf{X}_{t-1} + \widehat{cday}_{t-1} + \xi_t. \tag{15}$$

Tables 10 and 11 present the results of the estimation of (14) and (15). Table 10 shows that the forecasting regression of real returns has no explanatory power in line with the results in Table 2. It can also be seen that asset returns are an important explanatory of financial wealth growth. Moreover, the housing wealth growth regression confirms the persistence of the variation in this component of asset wealth: the lagged housing wealth growth is highly significant and this equation explains 71% of the variation in housing wealth.

Table 10: Estimates from Vector-autoregressions (VAR) specified in equation (14).

	Equation				
Dependent variable	r_t	Δc_t	Δf_t	Δu_t	Δy_t
r_{t-1}	-0.108	0.015	0.525*	0.019	0.029**
(t-stat)	(-1.034)	(1.587)	(16.325)	(1.206)	(2.175)
Δc_{t-1}	-1.218	-0.261**	0.625***	0.391**	-0.261***
(t-stat)	(-1.005)	(-2.424)	(1.675)	(2.092)	(-1.695)
Δf_{t-1}	-0.218	0.008	-0.056	-0.005	-0.000
(t-stat)	(-1.215)	(0.473)	(-1.017)	(-0.174)	(-0.015)
Δu_{t-1}	0.240	0.055***	-0.081	0.781*	0.142*
(t-stat)	(0.634)	(1.639)	(-0.696)	(13.408)	(2.963)
Δy_{t-1}	-0.528	0.152**	0.487**	0.240***	-0.062
(t-stat)	(-0.663)	(2.138)	(1.984)	(1.957)	(-0.616)
θ	0.037*	0.005*	-0.001	-0.001	0.005*
(t-stat)	(3.053)	(5.048)	(-0.268)	(-0.617)	(3.582)
$\bar{R}^2$	0.000	0.075	0.751	0.713	0.090

This table reports the estimated coefficients from vector-autoregressions (VAR).

Symbols *, **, *** represent, respectively, significance level of 1%, 5% and 10%.

Newey-West (1987) corrected t-statistics appear in parenthesis. The sample period is 1977:Q4 to 2001:Q1.

[30] The selected optimal lag length is 1, in accordance with findings from Akaike and Schwarz tests. However, the results are not sensible to different lag lengths.

Table 11 provides very similar results. The most important feature of this estimation is that it shows that *cday* conveys important important for forecasting asset returns: the $\bar{R}^2$ reaches 0.05 and the coefficient associated to *cday* is significant and important in magnitude (1.898) in line with the results of Table 2.

Table 11: Estimates from Vector-autoregressions (VAR) specified in equation (15).

	Equation				
Dependent variable	r_t	Δc_t	Δf_t	Δu_t	Δy_t
r_{t-1}	-0.187***	0.020**	0.514*	0.022	0.018
(t-stat)	(-1.740)	(2.099)	(15.199)	(1.277)	(1.324)
Δc_{t-1}	-2.022***	-0.205***	0.513	0.414**	-0.373**
(t-stat)	(-1.644)	(-1.851)	(1.322)	(2.122)	(-2.405)
Δf_{t-1}	-0.072	-0.003	-0.036	-0.009	0.020
(t-stat)	(-0.388)	(-0.161)	(-0.613)	(-0.310)	(0.855)
Δu_{t-1}	0.473	0.039	-0.049	0.775*	0.175*
(t-stat)	(1.238)	(1.129)	(-0.404)	(12.779)	(3.629)
Δy_{t-1}	-0.219	0.130***	0.530**	0.231***	-0.019
(t-stat)	(-0.278)	(1.831)	(2.130)	(1.848)	(-0.194)
θ	-2.554**	0.187***	-0.362	0.076	-0.356*
(t-stat)	(-2.323)	(1.891)	(-1.042)	(0.433)	(-2.567)
$\widehat{cday}_{t-1}$	1.898**	-0.133***	0.264	-0.056	0.265*
(t-stat)	(2.356)	(-1.837)	(1.039)	(-0.440)	(2.607)
$\bar{R}^2$	0.049	0.099	0.752	0.711	0.146

This table reports the estimated coefficients from vector-autoregressions (VAR).

Symbols *, **, *** represent, respectively, significance level of 1%, 5% and 10%.

Newey-West (1987) corrected t-statistics appear in parenthesis. The sample period is 1977:Q4 to 2001:Q1.

Using the VAR estimated in Table 10,[31] I also assess the change in expected future returns and consumption growth caused by a shock to any of the forecasting variables considered. Figures 8 and 9 report, respectively, the impulse-response functions of quarterly real returns and consumption growth to a one standard deviation impulse in each of the regressors. Figure 8 shows that while financial wealth shocks have a positive effect on real returns, housing wealth shocks have a negative effect on

[31] The same results are obtained using the VAR estimated in Table 11 and specified in equation (15).

real returns. Figure 9 shows that housing wealth shocks have a very persistent effect on consumption, while financial wealth shocks are mainly transitory.

Figure 8: Impulse-Response functions of real returns.

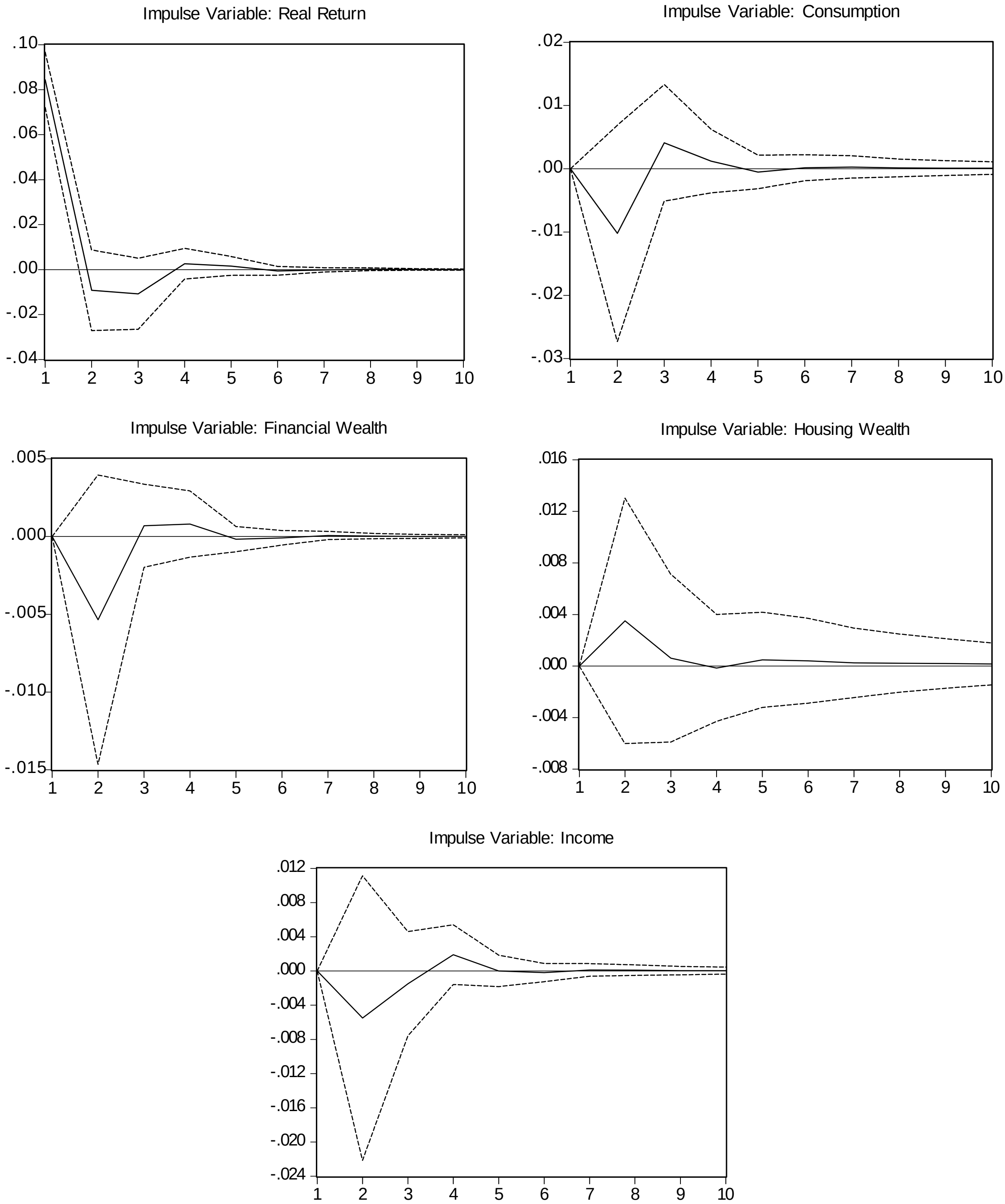
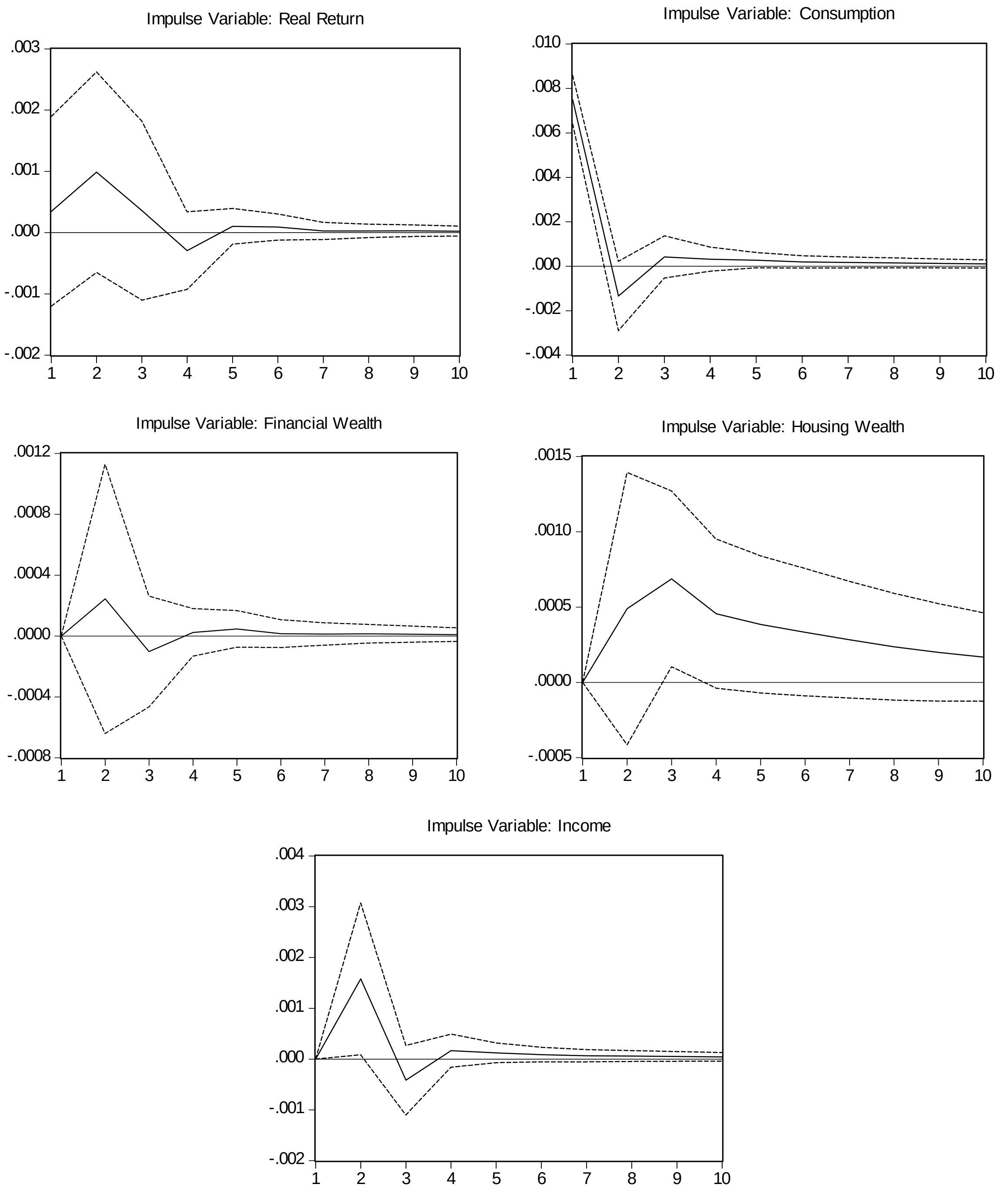


Figure 9: Impulse-response funtions of consumption.



5 Conclusions

This paper uses the representative consumer's budget constraint to derive an equilibrium relation between the trend deviations among consumption, (dis)aggregate wealth and labor income (summarized by the variable *cday*) and expected future asset returns, and explores predictive power of the empirical counterpart of these trend deviations ($\widehat{cday}$) for future asset returns.

The main finding of the paper is that $\widehat{cday}$ has high predictive power for future market returns and it performs better than a variable like $\widehat{cay}$ suggested by Lettau and Ludvigson (2001), which does not take into account the issue of the disaggregation of wealth. I show that the superior forecasting power of *cday* is due to: (*i*) its ability to track the changes in the composition of asset wealth and the specificities of the different assets; and (*ii*) the faster rate of convergence of the coefficients to the "long-run equilibrium" parameters. Therefore, disaggregating asset wealth into its main components (financial and housing wealth) is important and helps providing better forecasts for future asset returns: if consumption is above its trend relationship, then agents expect higher financial asset returns or higher housing asset returns.

Using data for the United Kingdom, I also show that financial wealth effects are significantly different from housing wealth effects. Unlike Lettau and Ludvigson (2001, 2004), who argue that asset wealth fluctuations are largely transitory, the results suggest that, while substancial fluctuations in financial assets need not indeed be associated with large subsequent movements in consumption, fluctuations in housing assets are very important due to their persistence. An important implication is that governments and central banks need to pay special attention to the behavior of housing markets (and to a smaller extent to the behavior of financial markets) when defining macroeconomic stabilizing policies.

This work is, however, only a first approach to the subject and has, therefore, some limitations. First, this approach does not correspond to a more structural representation of the economy in which the consumer's preferences and the production side are formalized. Lantz and Sartre (2001) address partially this question, showing that consumption does not react directly to wealth changes, but instead both consumption and wealth react to changes in productivity. Second, the formulation ignores labor income risk and its importance in the context of forecasting asset returns, an issue which has been dealt recently by Julliard (2004). Third, the specification implicitly assumes that agents consume a single good. In contrast, Lustig and Van Nieuwervurgh (2004) present a model in which agents care about the composition of a consumption basket that includes housing services.

Finally, this work is just a starting point for future research. A potentiality to analyze in the future is the role of financial deregulation/liberalization. Bayoumi (1993) and Caporale and Williams (1997), among others, point out the importance of these processes for the credit expansion and the elimination of liquidity restrictions that they provide; Bonser-Neal and Dewenter (1999) emphasize the effects of level of development of financial markets on the savings rate; and Bekaert *et al.* (2001) emphasize their importance for economic growth. Therefore, it would be important to approach the importance of these processes on the magnitude of wealth effects, an aspect that is analyzed in a recent work of Boone *et al.* (2001) and what are their implications for forecasting asset returns. Second, it would be also important to analyze the importance of the concentrated nature of the wealth and its impact on the dynamics of wealth distribution. Finally, although literature emphasizes the role played by wealth on non-durable consumption expenditure, it would also be interesting to analyze its role on durables consumption expenditure.

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Appendix

A Data Description

Consumption

Consumption is defined as total consumption (ZAKV) less consumption of durable (UTIB) and semi-durable goods (UTIR). Data are quarterly, seasonally adjusted at an annual rate, measured in millions of pounds (2001 prices), in per capita and expressed in the logarithmic form. Series comprises the period 1963:Q1 - 2003:Q4. The source is Office for National Statistics (ONS).

Wealth

Aggregate wealth is defined as the net worth of households and nonprofit organizations, this is, the sum of financial wealth and housing wealth. Data are quarterly, seasonally adjusted at an annual rate, measured in millions of pounds (2001 prices), in per capita terms and expressed in the logarithmic form. Series comprises the period 1975:Q1 - 2004:Q1. The sources of information are: Fernandez-Corugedo *et al.* (2003), for the period 1975:Q1 - 1986:Q4; Office for National Statistics (ONS), for the period 1987:Q1 - 2004:Q1.

Financial wealth

Financial wealth is defined as the net financial wealth of households and nonprofit organizations (NZEA). Data are quarterly, seasonally adjusted at an annual rate, measured in millions of pounds (2001 prices), in per capita terms and expressed in the logarithmic form. Series comprises the period 1970:Q1 - 2004:Q1. The sources of information are: Fernandez-Corugedo *et al.* (2003), for the period 1970:Q1 - 1986:Q4; Office for National Statistics (ONS), for the period 1987:Q1 - 2004:Q1.

Housing wealth

Housing wealth is defined as the housing wealth of households and nonprofit organizations and is computed as the sum of tangible assets in the form of residential buildings adjusted by changes in house prices (CGRI), the dwellings (of private sector) of gross fixed capital formation (GGAG) and Council house sales (CTCS). Original data is annual. Quarterly data was interpolated from original data using the following methodology: at the end of each year, any difference between Housing Wealth and CGRI is split into four and evenly distributed over the four quarters in that year (e.g. one quarter of the difference is put in the first quarter of the year, half in the second quarter, three quarters in the third, and the full difference in the fourth). Data are, therefore, quarterly, seasonally adjusted at an annual rate, measured in millions of pounds (2001 prices), in per capita terms and expressed in the logarithmic form. Series comprises the period 1975:Q1 - 2004:Q1. The sources of information are: Fernandez-Corugedo *et al.* (2003), for the period 1975:Q1 - 1986:Q4; Office for National Statistics (ONS), for the period 1987:Q1 - 2004:Q1. For data on house prices, the sources of information are: Office of the Deputy Prime Minister (ODPM), Halifax Plc and the Nationwide Building Society.

After-tax labor income

After-tax labor income is defined as the sum of wages and salaries (ROYJ), social benefits (GZVX), self employment (ROYH), other benefits (RPQK + RPHS + RPHT - ROYS - GZVX + AIIV), employers social contributions (ROYK) less social contributions (AIIV) and taxes. Taxes are defined as (taxes on income (RPHS) and other taxes (RPHT)) x ((wages and salaries (ROYJ) + self employment (ROYH)) / (wages and salaries (ROYJ) + self employment (ROYH) + other income (ROYL - ROYT + NRJN - ROYH)). Data are quarterly, measured in millions of pounds (2001 prices), in per capita terms and expressed in the logarithmic form. Series comprises the period 1974:Q3 - 2003:Q4. The sources of information are: Fernandez-Corugedo *et al.* (2003), for the period 1974:Q3 - 1986:Q4; Office for National Statistics (ONS), for the period 1987:Q1 - 2003:Q4.

Population

Population is defined as mid-year estimates of resident population of the United Kingdom (DYAY) in millions. Original data are available as an annual series. The data are interpolated to quarterly frequencies, computing the annual population growth rate and the applying the average quarterly population growth rate every quarter. Series comprises the period 1946:Q4 - 2003:Q4. The source of information is Office for National Statistics (ONS).

Price deflator

The nominal consumption, wealth, financial wealth, housing wealth, labor income and interest rates were deflated by the All Items-Retail Prices Index (CHAW) (January 13 1987 = 100). Data are quarterly. Series comprises the period 1947:Q4 - 2004:Q4. The source of information is Office for National Statistics (ONS).

Inflation rate

Inflation rate was computed from price deflator. Data are quarterly. Series comprises the period 1947:Q3 - 2004:Q4. The source of information is Office for National Statistics (ONS).

Interest rate ("Risk-free rate")

Risk-free rate is defined as the quarterly real yield rate of 3-month Treasury Bills (AJRP). Original data are available as an annual series. Quarterly data are computed applying the average quarterly real yield rate every quarter. Series comprises the period 1972:Q1 - 2004:Q4. The source of information is Office for National Statistics (ONS).

Asset returns

Asset returns were computed using the MSCI - UK Total Return Index for the UK, which measure the market performance, including price performance and income from dividend payments. We use the index which includes gross dividends, this is, approximating the maximum possible dividend reinvestment. The amount reinvested is the dividend distributed to individuals resident in the country of the company, but does not include tax credits. Series comprises the period 1970:Q1 - 2004:Q4. The source of information is Morgan Stanley Capital International (MSCI).

B Tests of the existence of unit roots and cointegration

Table B1: Phillips-Perron Unit Root Tests to the variables' cointegration order (variables in levels).

		Critical values	
	Phillips-Perron t-Statistic	5% Level	1% Level
c_t	-1.83	-3.46	-4.06
f_t	-3.03	-3.46	-4.06
u_t	-1.76	-3.46	-4.06
y_t	-2.60	-3.46	-4.06

Symbols * and ** denote rejection of the null hypothesis at a significance level of 1 and 5%, respectively.

Newey-West (1987) bandwith selection. Critical values suggested by MacKinnon (1996).

Table B2: Phillips-Perron Unit Root Tests to the variables' cointegration order (variables in first-order differences).

		Critical values	
	Phillips-Perron t-Statistic	5% Level	1% Level
Δc_t	-11.06*	-2.89	-3.50
Δf_t	-10.81*	-2.89	-3.50
Δu_t	-3.09**	-2.89	-3.50
Δy_t	-10.41*	-2.89	-3.50

Symbols * and ** denote rejection of the null hypothesis at a significance level of 1 and 5%, respectively.

Newey-West (1987) bandwith selection. Critical values suggested by MacKinnon (1996).

Table B3: Test of cointegration using the methodology of Engle and Granger (1987).

		Critical values	
	Phillips-Perron t-Statistic	5% Level	1% Level
$\widehat{cday}_t$	-5.47*	-2.89	-3.50
$\widehat{cay}_t$	-4.50*	-2.89	-3.50

Symbols * and ** denote rejection of the null hypothesis at a significance level of 1 and 5%, respectively.

Newey-West (1987) bandwith selection. Critical values suggested by MacKinnon (1996).