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Aggregate Consumption and the Stock Market: Should We Worry About Non-linear Wealth Effects?

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Aggregate consumption and the stock market: should we worry about non-linear wealth effects?

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Abstract: The linkage between stock market and aggregate consumption has been extensively studied in the context of linear econometric models. This paper proposes a less restrictive approach: short-run dynamics in US consumption are analysed applying semi-parametric techniques to a large sample of monthly data (1967-2002). This allows a rigorous assessment of the claim that consumers react differently to negative and positive changes in the value of their portfolios, or that they are only sensitive to "large" equity price corrections. The data display indeed non-linearities of this type, but their significance is modest; the results corroborate the traditional view that, overall, Wall Street is not a major concern for American households.

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1. Introduction

Is there a significant short-run relationship between stock market wealth and consumption expenditure at the aggregate level? What makes this question difficult to answer is the *short-run* and *aggregate* nature of it. If analysed in long-run terms for a single household owning equities, the issue is not very interesting; elementary budget-constraint algebra suggests that, insofar as it increases household's wealth, any portfolio appreciation will stimulate a rise in future consumption. Aggregation complicates things for mainly one reason: even in economies where the participation rate into the stock market is large, the distribution of asset holdings among families is very uneven. Clearly the budget constraint logic has to hold in the aggregate as well, but its empirical relevance will depend on the average ratio of stock wealth to household's global net worth and possibly on several, more complex features of the distribution. Furthermore, the timing of the adjustment is not trivial, because it is potentially related to a host of different factors –investor's perception of the temporary/permanent nature of the change in net worth, liquidity of the assets, nature of the fiscal system, etc.

Muddying the waters further, some authors recently suggested the possibility that agents react differently to positive and negative wealth changes, or that only large changes impinge on consumption choices. Since short-run wealth changes are dominated by equity price movements, a non-linear "wealth effect" would create a highly destabilizing link between stock market fluctuations and real aggregate activity. On the other hand, there are episodes where consumption kept growing at a steady pace in the face of large equity prices corrections: the 1987 US crisis is a frequently cited example. Many good reasons motivate a careful analysis of these issues. Monetary policy is one of these, especially in the light of the recent debate about the opportunity for central banks to "lean against" stock market misalignments. This paper investigates the nature of short-run aggregate equity wealth effects with a sample of monthly observations describing the US economy over the 1967-2002 period. The monthly frequency is somehow a novelty, as only quarterly or yearly data have been used in the literature. The rationale behind this choice is twofold. Firstly, equity prices fluctuate substantially over very short horizons; equities represent indeed the most volatile component of households' wealth. In this sense, a good deal of information is lost when looking at quarterly changes, and a finer time grid allows a better analysis of short-run issues. Secondly, the large number of observations makes it possible to "take seriously" the possibility of a non-linear linkage between consumption and the stock market; in particular, it permits a test of ordinary linear equations against an unrestricted semi-parametric model. The philosophy of this paper is to (i) aim at the maximum possible generality and (ii) impose a minimal amount of structure on the data. Hence, rather than postulating a complete model, the analysis relies on a simple budget constraint argument that holds under a broad range of optimal consumption theories. Local polynomial estimation, then, makes it possible to relax the assumption that consumption and wealth are linearly linked without *a priori* formulating any specific alternative hypothesis. The data display some non-linearity: losses count more than gains, and small changes in equity prices tend not to count at all. The main conclusion of the paper, though, is that these phenomena are hardly significant from the statistical point of view and the linear model provides overall a fairly good description of the data.

Section 2 reviews the literature focussing on the empirical work closer to this paper. Section 3 discusses existing evidence on non-linear equity wealth effects and presents *prima facie* results from a univariate non-parametric analysis. Section 4 describes the data to be used in the multivariate set-up (two alternative measures of consumption, labour income, equity wealth, non-equity wealth, interest rates). Section 5 analyses the long-run properties of the variables in the context of a VAR-VECM model. In section 6, the short-run consumption equation implied by the VECM is compared to its semi-parametric counterpart; all technical details on the latter are discussed in the appendix. Section 7 concludes[*].

2. Consumption and wealth: an overview.

Most of the recent empirical work on aggregate wealth effects in the US applies cointegration and error-correction methodology to quarterly data. The existence of a common trend between aggregate consumption, labour income and wealth is indeed predicted by representative consumer models as well as life-cycle models on the basis of the "budget constraint algebra" mentioned in the introduction (Campbell and Mankiw 1989; Lettau and Ludvigson 2001, 2003)[1]. Following Lettau and Ludvigson (2001), consider a representative agent who earns a net return R_{t+1} on his period-t wealth W_t (inclusive of both financial and human assets); the accumulation process is described by:

$$W_{t+1} = (1 + R_{t+1})(W_t - C_t) \qquad [1]$$

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or equivalently $\log(W_{t+1}) = r_{t+1} + \log(W_t - C_t)$, where $r_t \equiv log(1+R_t)$. Under the only additional assumption of the C/W ratio being stationary, a first order Taylor approximation of the log equation yields (ignoring constants)[2]:

$$\Delta w_{t+1} \approx r_{t+1} + (1 - 1/\rho)(c_t - w_t) \qquad [2]$$

where $\rho \equiv (W\text{-}C)/W$ is the steady-state investment-to-wealth ratio. By solving the equation forward and imposing $\lim_{i\to\infty} \rho^i (c_{t+i} - w_{t+i}) = 0$, the following expression for the consumption-wealth ratio obtains:

$$c_t - w_t = \sum_{i=1}^{\infty} \rho^i \left(r_{t+i} - \Delta c_{t+i} \right). \qquad [3]$$

This equation is directly implied by the intertemporal budget constraint, it does not depend on any specific assumption on preferences, and it holds both *ex post* and *ex ante* – so that c_t-w_t also equals the expectation of the right hand side conditional on time-t information. Aggregate wealth is defined as the sum of non-human and human assets: $W_t = A_t + H_t$. A logarithmic approximation of the equation delivers:

$$w_t \approx \omega a_t + (1 - \omega) h_t \qquad [4]$$

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[1] King et al. (1988) provides a survey of representative consumer models developed in the real business-cycle literature; Galì (1990) obtains the common trend in the context of a life-cycle model.

[2] The equation is derived in the appendix of Campbell and Mankiw (1989).

where ω≡A/W is the steady-state value of the assets-wealth ratio. A common strategy to get around the non-observability of H consists of assuming that its non-stationary component is related to labour income Y so that $h_t = b + y_t + z_t$, where z_t is zero-mean, stationary random variable. Substituting [4] and the equation for h in [3] gives:

$$c_t - \omega a_t - (1-\omega) y_t = E_t \sum_{i=1}^{\infty} \rho^i \left(r_{t+i} - \Delta c_{t+i} \right) + (1-\omega) z_t . \qquad [5]$$

The presumed stationarity of the terms on the right-hand side implies that c, a and y cointegrate. As for a, the theory does not rationalise different roles for different assets: why should equities deserve a special treatment? Several arguments have been put forth (usually in an informal way) suggesting that the "stock market wealth effect" has somewhat peculiar features. Poterba (2000) notes that the distribution of equities among households is highly concentrated, and equities are usually kept in tax-favoured retirement accounts and considered "long term assets" by consumers. As a consequence, the propensity to consume out of stock market wealth might be smaller than the MPC out of net worth; furthermore, due to the growing importance of equities in household wealth, the overall MPC may have declined in the nineties. These conjectures are broadly supported by Mehra's (2001) empirical results. The author uses a sample spanning from 1959.Q1 to 2000.Q2 and considers alternative measures for consumption (per capita real consumer spending, per capita real consumer spending on non-durable goods and services) and wealth (global net worth, equity wealth). In all cases he finds evidence of a single cointegrating relationship between consumption, income and wealth. The elasticity of consumption is indeed smaller with respect to equity wealth (.02) than with respect to non-equity wealth (.15), even though the implied level responses are similar and roughly equal to .03. By replicating the analysis on sub-samples 1960.Q2-1990.Q2 and 1960.Q2-1995.Q2, Mehra also finds that the propensity to consume out of total wealth declined during the nineties whereas the propensity to consume out of equity wealth remained constant, which is consistent with a changing composition of wealth stocks. In the short-term equation, contemporaneous equity and non-equity wealth (instrumented by their own lags) are both significant, with coefficients .01 and .10. Simple computations show that – even with such small estimates – the stock market boom likely added an average 1% to the annual GDP growth rate throughout the 95-99 period.

Ludvigson and Steindel (1999) use quarterly data from 1953 to 1997 and focus on total net worth, pointing out that quarterly fluctuations in net worth are in any case driven by its financial component. The outcome of the estimation procedure is a VECM in log-differences of consumption, wealth and labour income. There is a unique cointegrating vector, and the implied MPC for variables in levels are 0.046 (wealth) and 0.718 (income). Interestingly, consumption growth predicts both wealth and income growth but neither of these predicts consumption; furthermore, the loading factors suggest that wealth does most of the adjustment to restore the long-run equilibrium. The authors perform impulse-response analysis under two different timing assumptions. If consumption is constrained to respond with a 1-quarter lag, the impact of a wealth shock is statistically negligible at all horizons; if contemporaneous adjustment is allowed, consumption jumps significantly on impact and the response is over by the end of the quarter[3].

[3] A short-lived movement in consumption growth obviously implies a permanently higher consumption level.

A thorough assessment of the correlation between wealth and consumption is provided by Lettau and Ludvigson (2003). On post-war quarterly data (c_t, w_t, y_t) cointegrate with coefficients (1,-0.30,-0.60), and only w_t adjusts to restore the long-run equilibrium. Relying on the VECM restrictions, the authors identify permanent and transitory innovations and perform a variance decomposition along the lines of King et al. (1991), Gonzalo and Granger (1995), Gonzalo and Ng (2001). The main result is that transitory innovations generate up to 88% of the variation in net worth, whereas variation in aggregate consumption is dominated by permanent shocks. Consumption responds differently to transitory and permanent changes in wealth, and temporary wealth fluctuations (the majority) are unrelated to consumption. The authors conclude that the "marginal propensity to consume" derived from common trend estimates, which measures the marginal impact of a permanent wealth change, overstates the magnitude of the channel linking consumption and wealth.

Wealth effects have also been studied on household level data. An example of this branch of literature is the attempt by Dynan and Maki (2000) to disentangle *direct* and *indirect* channels by using data from the *Consumer Expenditure Survey* (1983.Q1 to 1999.Q1)[4]. The authors consider a simple equation in log-differences relating consumption of non-durables and services to lags of the Wilshire 5000 index and control variables; the equation is estimated for stockholders and non-stockholders separately, alternatively classifying as "stockholders" households with security holdings greater than \$0, \$1,000 or \$10,000. A 5%-significant positive correlation between consumption and lags of the price index is documented for stockholders only. According to the authors, the evidence suggests at the same time a relevant *direct effect* and a negligible *indirect effect*. This literature can also be linked to the *uncertainty hypothesis* formulated by Romer (1990) with reference to the 1929 crisis[5]. If a negative equity price correction signals (or generates) uncertainty about the future, it may induce consumers to postpone or reduce purchases of durable goods. Hence, the observed change in non-durable expenditure may result from a negative *income effect* and a positive *substitution effect*.

There are only a few empirical studies on countries other than the US (a recent exception is Bertaut, 2002). This depends on data availability, but also on the shared belief that the US are the country where wealth effects (and particularly stock market wealth effects) can be studied most effectively. Three main factors support this conclusion. Firstly, by the end of the run-up in equity prices in 1999, the estimated ratio of outstanding equities to GDP was over 180% for the US and the UK, and below 100% for France, Germany, Japan (Bertaut, 2002). Secondly, according to the *Survey of Consumer Finances*, 49% of US households held equities in 1998, either directly or through mutual funds and retirement accounts; analogous national surveys suggest figures of 37% for Canada and 27% for the UK, with the big economies in continental Europe all lying below 15%. Finally, stock ownership appears to be more concentrated in the EU than in the US (IMF, May 2000 *World Economic Outlook*).

[4] In the authors' terminology, the "direct" channel links consumption to wealth through the budget constraint; the "indirect" channel encompasses all complementary (or alternative) linkages, such as the possibility that stock prices simply predict future changes in consumption.

[5] The hypothesis stresses the difference between durable and non-durable goods. Since durable goods generate some lock-in effect, consumers might prefer not to buy them when uncertainty about their future income stream is high. By discouraging consumption of durables, a rise in uncertainty increases the resources available for non-durable shopping; if durables and non-durables are to some extent substitutes, this translates into an increase in non-durables consumption. Romer (1990) shows that data on retail sales fit these predictions.

3. The case for non-linearity: a first non-parametric glance at aggregate consumption.

The first evidence on possibly non-linear wealth effects emerged from the research on lotteries and bequests (Poterba, 2000). Small lottery winnings (less than $15,000) seem to have no discernible impact on households' behaviour, whereas large winnings induce an increase in some combination of spending and leisure time. Insofar as they induce genuine exogenous shifts in net worth, lottery prizes and stock market gains should have the same effects. Furthermore, when the source of wealth shocks is the stock market, an informational issue may arise: small fluctuations in price indices usually do not receive attention from the media, so that an average (non professional) investor is unlikely to be fully aware of his "small" gains and losses[6].

Shirvani and Wilbratte (2000) report that after his speech at the New York Economic Club in December 1997, Alan Greenspan suggested more econometric work should address the nature of stock market wealth effect, particularly in the light of the potentially asymmetric impact of positive and negative equity price fluctuations. In their view, three factors may generate the asymmetry. The first one is the convexity of consumers' utility under risk aversion. If an agent's utility function is convex in cash-on-hand, a negative wealth shock determines a larger absolute change in the utility than an equally big positive shock. Hence, the agent is more willing to cut consumption and recover the optimal cash-on-hand level after the loss than to increase consumption in the opposite case. The second factor has to do with the fiscal system. When stock prices

[6] The idea that individuals are more sensitive to losses than gains has been the basis for several departures from the standard Von Neumann-Morgenstern approach to modelling agents' preferences, including *habit formation* (e.g. Constantinides, 1990) and *loss aversion* (e.g. Kahneman and Twersky, 1979, 1992; Benartzi and Thaler, 1995), which are based on different behavioural intuitions but share the relaxation of the time-separability assumption. A discussion of these theories is beyond the scope of the

change, the implied change in tax liabilities drives a wedge between the market value of the stocks and their net value for the owner. If capital gains are taxed progressively and capital losses are not entirely deductible, a depreciation impinges more heavily on net wealth than an appreciation. Finally, liquidity constraints may play a role: consuming less is always possible for everybody, whereas some consumers may find it difficult to borrow and increase their consumption level after an increase in their net worth. Hence, it is possible that aggregate consumption is relatively more sensitive to stock market downturns. In order to test this proposition, the authors consider 1970:Q1-1996:Q2 data on aggregate consumption (C), national income (Y), M2 money supply (M) and a stock price index (S) for the US, Germany and Japan; when formulating the VECM, they use a dummy variable to separate positive and negative equity price changes. For each country, they estimate the following consumption equation:

$$\Delta C_t = \mu + \sum_j \phi_j ec_{j_{t-1}} + \sum_i \beta_i \Delta C_{t-i} + \sum_i \gamma_i \Delta Y_{t-i} + \sum_i \delta_i \Delta M_{t-i} + \sum_i \theta_i^+ \Delta S_{t-i}^+ + \sum_i \theta_i^- \Delta S_{t-i}^- + \nu_t \qquad [6]$$

where the ec_j are cointegration residuals. The θ^+_i and the θ^-_i are jointly significant for every country; the difference $(\Sigma\theta^+_i - \Sigma\theta^-_i)$ is significant at the 5% level for the US and at the 10% level for Germany and Japan, and it is negative in all three cases, consistently with the arguments above. The paper, though, is questionable in many respects. All forms of wealth other than equities are ignored; furthermore, the closed-economy assumption implicit in the choice of variables makes sense for the US but seems hard to defend in the case of Germany and Japan. Bertaut (2002) follows the same strategy of Shirvani and Wilbratte (2001) using quarterly data on several countries from 1981 to

paper, but of course they provide natural frameworks for the interpretation of any non-linearity found in consumption data.

1998. The paper estimates a range of models; in one of them consumption growth rate is explained by income, stock prices, a short-term interest rate and the unemployment rate. When stock price changes are split depending on their sign, the asymmetry appears to be country-specific: in Japan consumption responds more significantly to equity devaluations, but in Canada and US it is the opposite. The author ascribes the conflict between her results for the US and Shirvani and Wilbratte's to generic specification and sample differences.

To summarise, there is some evidence that consumption may adjust non-linearly to movements in equity prices, but the empirical literature is at a very early stage and it has not conveyed a clear message yet. Semi-parametric techniques present a new perspective, because they allow a completely unrestricted analysis of the adjustment process. How does $E_t\Delta c_{t+1}$ look like if we only condition on stock market wealth? Figure 4 presents a local polynomial estimation of the following model:

$$g^c_t = f(g^{we}_{t-i}) + \varepsilon_t \qquad [7]$$

where g^c_t is the yearly growth rate of non-durable consumption and g^{we}_{t-i} is the yearly growth rate of the Standard & Poors 500 index at lags of 1, 2, 3 and 6 months[7]. Linearity is clearly an appropriate assumption in the case of g^{we}_{t-6} (bottom right graph); in fact, all wealth lags beyond the 4th generate a neat linear function. But as the time gap is reduced, the straight line progressively turns into a more interesting function (e.g. top left graph). The reaction to a recent decrease in equity value $(g^{we}_{t-1}<0)$ is concave: the larger g^{we}_{t-1} (in absolute value), the larger the expected fall in consumption. Small, positive increases in equity values correspond to a noisy area where consumption seems not to respond at all. Finally, large equity gains *($0.2<g^{we}_{t-1}$)* stimulate consumption to some extent, but "booms" are clearly less powerful than "recessions" – the right branch of the curve is overall flatter than the left one.

The model is clearly oversimplified[8]; what is interesting, though, is that the non-parametric estimates seem to capture in a remarkably consistent fashion three facts highlighted by previous works: an asymmetry between positive and negative wealth changes, a "lottery effect" by which large changes count relatively more, and perhaps a qualitative difference between permanent and temporary shocks. Indeed, a possible interpretation for the fact that the non-linearity is only associated with recent price changes is that, as time elapses, the transitory nature of most price fluctuations is revealed. The shape of the mean function is similar for total consumption, even though the non-linear features are generally less pronounced, and it is robust to the choice of the stock price index (replacing S&P 500 with Wilshire 5000 does not change the outcome).

A simple way to gain some insight on the significance of the non-linear features consists of parameterising f as a piece-wise linear function with an unknown threshold[9]:

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[7] The estimation is a *LOESS*, namely a local linear polynomial fit based on the tricube kernel. The choice of the bandwidths is arbitrary at this stage, but the estimates in section 6 will be based on an asymptotically optimal bandwidth selected by cross-validation (see appendix).

[8] If aggregate consumption were very close to a random walk the misspecification would not be too serious, but as a matter of fact the equation does not generate white noise residuals.

[9] Since there is an interval of values for g^{we} where the response of consumption is almost undetectable, one may argue that an equation with two thresholds is more appropriate:
$g^{c}_{t} = \alpha + \beta\, g^{we}_{t-1} + \delta_1 d_1(g^{we}_{t-1}-\zeta) + \delta_2 d_2(g^{we}_{t-1}-\xi) + \varepsilon_t$,
where $d_1 = I\{g^{we}_{t-1}\geq\zeta\}$ and $d_2 = I\{g^{we}_{t-1}\geq\xi\}$, $\zeta<\xi$ and I is an indicator function. This equation and equation [8] are non-nested insofar as the thresholds do not match (i.e. $\theta\neq\zeta$ and $\theta\neq\xi$). The double-threshold equation performs slightly better in terms of Schwartz information criterion, but the two estimated thresholds appear to be very close (-.01, -.009), suggesting that the model might be over-parameterised and that a simple single–threshold equation is preferable.

$$g^{c}_{t} = \alpha + \beta g^{we}_{t-1} + \delta\, d(g^{we}_{t-1}-\theta) + \varepsilon_t \qquad [8]$$

where $d=I\{g^{we}_{t-1}\geq\theta\}$ and I denotes an indicator function. The dummy variable is defined in such a way to impose continuity in θ; parameters and threshold can be jointly estimated in a maximum likelihood framework assuming normally distributed errors. The threshold is in $\theta=-.15$; the two elasticities are roughly 8% $(g^{we}_{t-1}<-.15)$ and 1% $(g^{we}_{t-1}>.15)$, with standard errors of .018 and .005. The null hypotheses that the slope is constant over the sample $(\delta=0)$ or changes depending on the sign of g^{we}_{t-1} $(\theta=0)$ are strongly rejected by the data[10]. At the very least, it is possible that ordinary linear models estimate a biased wealth effect resulting from the averaging of a "bad times" elasticity and a "normal times" elasticity. On the other hand, the presumption that the elasticity changes depending on the direction of the stock market movement may be wrong, as the 8% elasticity only comes into play for *large* negative price corrections. The robustness and significance of these features in a multivariate context will be assessed in section 6, after analysing the equilibrium behaviour of the variables and deriving a benchmark linear model.

4. Definitions and data issues.

Consumption and labour income are defined in a completely canonical way. The theory applies to consumption *flow*, which includes expenditure on non-durable goods and services and use of existing durable goods. This variable is measured as personal consumption expenditure on non-durable goods and services excluding shoes and clothing, available from the Bureau of Economic Analysis *(c^{nd})*. The implicit – and usual – assumption is that the true flow, which includes non-observable services extracted from durable goods, is a multiple of its observable part. I also report some estimates obtained using total consumption *(c)*, but it is worth stressing that the theory above gives no guidance on the equilibrium behaviour of this variable. Disposable labour income *(y)* is computed as in Ludvigson and Steindel (1999), and the source is again the BEA.

The choice to use monthly data comes at a cost, as the BEA *net worth* series used in most of the literature is only available at a quarterly frequency: equity wealth *(w^e)* is replaced by a proxy (the *Standard&Poors 500* price index), and the non-equity wealth series *(w^{ne})* is obtained by interpolating quarterly data. A brief discussion of this strategy is warranted. To the extent it measures the market value of capital, the price index could be considered a proxy for the agent's total wealth rather than his financial wealth. Figure 1 displays the quarterly pattern of two broad market indices (*Standard&Poors 500*, *Wilshire 5000*) together with a variety of wealth measures. The indices mimic quite closely net worth, but the fit progressively improves when the definition of wealth is restricted to financial assets and then corporate equities. In particular, the indices closely follow the equity wealth pattern in troubled periods such as 1973-75, 1987-88 and 2001. Figures 2 and 3 make the point more compelling showing the series in log-differences: the price index provides a fairly good description for equities, whereas total net worth is far "too smooth". In consideration of this, the *S&P500* price index (from Datastream) is scaled up to match the sample mean of *corporate equity holdings* (from the Flow of Accounts data) and used as a proxy for this specific item.

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[10] In the first test, a potential problem is that the threshold θ is not identified under the null H_0: $\delta=0$.

Obviously, there are many net worth items other than *corporate equities* whose value depends on the stock market (mutual fund shares, pension fund reserves, bank personal trusts). The inclusion of these items in "equity wealth" or "non-equity wealth" variables is to some extent arbitrary. The reason why they are considered "non-equity wealth" in this context is twofold. Firstly, these assets are held under a variety of contracts that drive a wedge between the crude market performance and the performance of households' portfolios. Secondly, a restrictive definition of equity wealth implies a relatively conservative approach, reducing the risk of overestimating the reactivity of consumption to stock market fluctuations. Non-equity wealth, namely net worth minus corporate equities, is relatively less volatile over short horizons, so an interpolation of quarterly data is likely to track the true monthly series reasonably well. Since net worth is measured in end-of-period value, it seems natural to assign each quarterly value to the last month of the quarter and place intermediate data points on the cubic spline. The Board of Governors of the Federal Reserve System supplies quarterly series for net worth and corporate equities. The estimation of wealth-effects also raises a timing issue (Lettau et al., 2001). Consumption is obviously measured as a (time-averaged) flow over a given period, whereas the net worth figure is a point-in-time estimate referring to the end of a period. Hence, the value of "time-t" wealth does not belong into the agent's information set when "time-t" consumption is decided. To by-pass the problem, the w^{ne} series is lagged once: the contemporaneous value is thus the stock cumulated by the end of the previous period.

The sample is 1967.01-2002.08. All nominal aggregate variables are deflated by the personal consumption expenditure chain-type price deflator with base year 1996 (from BEA) and divided by a population measure (from the FRED database maintained by the Federal Reserve Bank of St. Louis). All variables are in logarithms. The series display a fairly regular path, so the (inevitable) assumption that there are no idiosyncratic trend breaks seems justified.

5. The long-run.

The data are clearly non-stationary. Formal tests suggest that all variables are integrated of order one, and all variables except equity wealth possess a linear deterministic trend (details are available upon request). The existence of equilibrium relationship(s) can thus be uncovered by testing for cointegration on the basis of the procedure suggested by Johansen (1988, 1991). This testing procedure presumes that the variables are linked in a vector autoregressive model (VAR) of some order. Non-durable consumption is considered first. In a comparison of VAR models containing up to 8 lags, Akaike's and Schwarz's information criteria respectively suggest a specification of order 4 and 2. As appendix I shows, there is strong evidence supporting the existence of a single cointegrating relationship between c^{nd}, y, w^e and w^{ne} in all VARs up to the 4^{th} order. A dynamic OLS equation (Stock and Watson, 1993) delivers the following estimates[11]:

$$\underset{}{c^{nd}_t} = -2.14 + \underset{(0.030)}{0.931y_t} + \underset{(0.002)}{0.042w^e_t} + \underset{(0.021)}{0.218w^{ne}_t} \qquad [9]$$

[11] These results have been obtained including eight leads and lags of differenced variables; the estimates are not sensitive to this choice. Johansen's full information maximum-likelihood technique yields analogous figures. Of course estimates of the common trend based on 35 years of data have to be interpreted with some caution.

(standard errors in brackets; lead and lag terms are ignored). All parameters are in the range of values found in the literature. The elasticity of consumption with respect to labour income *(β_y)* is close to unity; Gali' (1990) interprets a similar result as evidence that life-cycle considerations are not decisive in determining consumption choices. The elasticity on equity wealth *(β_{we})* is much smaller than that on non-equity wealth *(β_{wne})*, as in Mehra (2001); however, using the average consumption-equity wealth ratio, the implied marginal impact of a dollar increase in equity wealth on consumption is about 5.2 cents, in line with common estimates of the MPC out of total wealth.

As mentioned in section 2, there are good reasons to believe that these parameters changed over time; in order to investigate this possibility, the dynamic OLS equation is estimated on a 10 years rolling sample (figure 4). The elasticity of consumption to equity wealth is quite stable over time, whereas there is evidence of substantial instability in β_{wne} and β_y; in particular, β_y (β_{wne}) is significantly larger (smaller) in the second half of the sample. An interesting feature of the figure is the symmetric pattern of the rolling estimates for these two parameters: starting from the mid seventies, there seems to be a substitution between assets and labour income as sources to fund consumption[12]. In the last 20 years the US experienced a decrease in the personal saving rate and, at the same time, an increase in the consumption/GDP ratio and the wealth/income ratio (Parker, 1999); the saving rate was actually almost zero in 2001. The fact that labour income is entirely spent on consumption fits well in this picture. Parker (1999) argues that intergenerational transfers by the Social Security system in recent years are favouring living generations, which under certainty increases agents' propensity to consume out of current income. Indeed, significant tax cuts were enacted in 2001 as well (Bertaut, 2002). The development of the financial system might have pushed in the same direction by relaxing liquidity constraints and perhaps reducing the precautionary motive for saving. Short-run dynamics can be described by a vector error correction restricted on the basis of the equilibrium relationship above:

$$\Delta x_t = \kappa + \gamma\beta' x_{t-1} + \Gamma(L)\Delta x_{t-1} + \varepsilon_t \, , \qquad [10]$$

where $x = (c^{nd}, y, w^e, w^{ne})$, $\beta'=(1, -\beta_y, -\beta_{we}, -\beta_{wne})$ and $\Gamma(L)$ is a lag polynomial. As table 1 shows, the error-correction term turns out to be significant only in the equation for Δw^e[13]. It is well established in the literature that in general it is wealth, and not consumption, that adjusts in the long-run; the fact that in particular *equity* wealth does all the adjustment might deserve further analysis, but is not examined in any greater detail here. Dynamics in Δc^{nd} can be described in terms of past consumption and – to some extent – labour income and non-equity wealth; none of the Δw^e lags is individually significant at the 10% level.

Total consumption behaves quite differently. In this case, the evidence on cointegration is ambiguous (see again appendix II). The tests yield different results depending on the VAR order; in a VAR(4) (selected by Akaike's criterion) the null hypothesis of no cointegration cannot be rejected at the 5% significance level, and in a VAR(2) (selected by Schwarz's criterion) the test statistics give conflicting indications. The "balance of evidence" suggests that equation [5] in section 2 breaks down when

[12] In Lettau and Ludvigson (1999), the cointegrating vector is estimated over three non-overlapping sub-samples and β_y displays an analogous increasing pattern.

[13] The table reports estimates for a VECM(3) because a Wald test does not reject the exclusion of the fourth lag of Δx_t while rejecting at the 1% the exclusion of the third lag. In terms of short-run analysis,

durable goods are included in the consumption measure. Accordingly, the short-run is modelled by a vector autoregression in first differences. Again, changes in equity wealth appear to be scarcely relevant independently of the order of the model, and consumption growth is strongly correlated with its own lags.

These estimates confirm the view that stock market and consumption are largely unrelated, especially in the short-run; impulse-response analyses (not reported) also show that c and c^{nd} do not respond significantly to equity price shocks. The issue is whether, given the results discussed in section 4, this conclusion depends on the linear models being inadequate.

6. A semi-parametric investigation of the short-run.

This section examines the potential gains from switching to a non-linear forecasting equation for consumption. Given that the data is monthly, Δc^{nd}_t and $(\Delta y_{t-1}, \Delta w^e_{t-1}, \Delta w^{ne}_{t-1})$ are "almost" contemporaneous; in particular, unless agents monitor their financial position on a weekly basis, Δw^e_{t-1} captures quite accurately the information on which time-t consumption decisions are based. Thus, if causality ran from wealth to consumption the error-correction equation could be regarded as a reduced-form consumption function. However, the direction of causality between wealth and consumption is a problematic issue on which this paper does not take a position. In this sense, a test of the null hypothesis that the forecasting equation is linear has to be interpreted with caution. If the null cannot be rejected (i.e. the linear equation passes the

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the crucial result is that there is no error-correction term in the consumption equation; again, this is test), we have a case against the existence of non-linear wealth effects; but if the null is rejected wealth effects are only one of the possible explanations, which is only legitimate if causality runs from wealth to consumption.

It is a common practise to expand consumption equations including variables that are likely to be influential in the short-run but irrelevant in terms of cointegration; the selection of lags can also be improved by focussing on equation-specific rather than global criteria. These issues are considered below before examining a semi-parametric formulation of the model, in order to minimize the probability of coming across "spurious" non-linearities generated by an incomplete specification of the linear equation. The literature presents a long list of variables that proved to be to some extent significant in consumption growth rate regressions; this paper focuses on real interest rates. In order to capture potential term structure effects, two different rates are considered: the one-month certificate of deposit (r^1) and the ten-years treasury bill (r^2)[14]. The selection of a parsimonious, well-behaving linear equation for Δc^{nd}_t is accomplished in a general-to-specific way, initially allowing for 12 lags of each regressor and progressively removing insignificant lags from the model[15]. With a monthly sample spanning more than 30 years, structural stability is also a concern. Structural breaks are identified by recursively estimating the equation with a single-observation rolling dummy variable; the exact definition of each break period is then worked out compounding statistics and history.

robust to the lag choice.

[14] Once these two extremes are taken into account, rates with intermediate maturity have no additional explicative power in the model. Both series appear to be stationary; nominal rates are deflated by expected inflation, namely fitted values from an autoregressive model for actual inflation (as in Maclem, 1994). The data source is the Board of Governors of the Federal Reserve System.

[15] Regressors that are not significant at the 5% level are removed, unless this implies a worsening of the Schwartz information criterion. This approach is admittedly a-theoretical, but it seems to be the only choice in this context.

Table 3 reports the estimates. Several critical phases in the macroeconomic history of the US are accurately mirrored in this monthly dataset. The first three dummy variables can be easily interpreted in terms of oil price changes and international instability. The Arab oil embargo on exports to the US starts in October '73 and, by December, oil price increase fourfold; the embargo ends in March '74 *(a)*. In December 1979 world oil prices reach their sample maximum after an increase decided by Saudi Arabia; the beginning of 1980 also sees the first major fight between Iran and Iraq *(b)*. The third dummy *(d)* covers the period between the invasion of Kuwait by Iraq (August '90) and the end of the "Desert Storm" campaign (February '91). Interestingly, the 1987 stock market crisis is not highlighted in the recursive estimation. The causes behind the break in the middle of 1992 *(d)* are not clear. The US experience social frictions – Los Angeles riots take place between April and May; the crisis of the ERM, culminating when Italy and the UK exit the system on the "Black Wednesday" 16/9/92, fosters instability in the financial markets; some uncertainty might also stem from the approaching elections, and in particular from the possibility of a fiscal reform (president Clinton's mandate begins in November). These events, though, do not seem to provide a satisfactory explanation for the negative shift. Finally, in September 2001 the equation drops by almost thrice the intercept. All coefficients have reasonable signs. The estimation also uncovers an interest rate spread term (r^{sp}), whose positive influence on consumption probably depends on the spread capturing the perceived conditions and prospects of the economy. There is one significant equity lag (Δw^{e}_{t-10}), and the elasticity of Δc_t with respect to this form of wealth is less than 2% even after taking into account the feedback generated by lags of consumption. Again, the result is in line with the VECM in section 5 and the literature in general. No sign of misspecification is detected by diagnostic tests, and the residuals appear to be normally distributed, uncorrelated and homoscedastic (all details are available upon request).

The non-linear counterpart of this model is described in table 3 and figures 6 and 7. Keeping everything else unchanged, equity wealth is now allowed to enter the equation via an unknown real function *f*, which is estimated by fitting local linear polynomials; details on the estimation technique can be found in appendix II. A comparison between tables 2 and 3 reveals that all coefficients preserve their sign and magnitude, with minor changes in the significance level; the overall fit of the model is basically unchanged, a first indication that the non-linear features are not prominent. Figure 6 shows estimates of *f* for different bandwidths. The straight line results from a bandwidth covering the whole sample; in this case the local-linear estimate coincides with the OLS regression line. The curve is generated by a "small" bandwidth (in the notation used in the appendix, $k<k^*$), and the dots represent the regression estimated at the sample points using a range of intermediate bandwidths. As the bandwidth is reduced, the most significant discrepancies between the curves arise at the boundaries of the dataset. The volatility of local estimators depends on the density of the regressor, so that the estimator is inevitably less accurate at the boundaries. The consequences are made explicit in figure 7, where the optimal $(k=k^*)$ estimate of *f*, together with a 95% confidence band, is plotted against the distribution of Δwe.

The basic conclusion to be drawn from the figures is that the linear and non-linear alternatives are virtually indistinguishable: in a point-wise comparison, the linear model is never rejected[16]. Independently of the significance issue, the features of *f* are incompatible with the hypothesis in Shirvani and Wilbratte (2000). The positive branch of the straight line passes the non-parametric examination better than the negative one, as only for $\Delta we_{t-10} > 0$ does the lower bound of the confidence region depart from 0. Furthermore, even the worst performances of the Standard & Poors 500 index (which has a monthly change of -10% or less three times over the sample, in 1974, 1987 and 2001) are perfectly compatible with a linear conditional mean function. The idea that "large" price fluctuations are somehow more influential is not supported by the data.

7. Conclusions

This paper analyses short-run dynamics in US aggregate consumption data, using a large sample of monthly observations to assess the nature and relevance of the *stock market wealth effect*. The strategy pursued consists of two steps. Firstly, long-run equilibria are investigated in the context of a VAR-VECM model in consumption, labour income, equity wealth (as measured by the Standard & Poors 500 index) and non-equity wealth. In the unique cointegrating vector, the elasticity of consumption to equity wealth is smaller than that to other forms of wealth; a $1 increase in equity value implies a 5 cents increase in consumption.

The VECM short-run consumption equation (expanded to include interest rates) is then compared to a *partial linear model* where no functional restriction is placed on the linkage between equity prices and consumption. The error-correction model suggests a short-run elasticity below 2%, and the semi-parametric investigation shows that there is no substantial evidence of non-linearity. In particular, the claim that the wealth channel

[16] The confidence band might seem surprisingly large, but this only depends on the scaling of the picture (10^{-3} for the vertical axis, 10^{-2} for the horizontal one). The standard regression also outperforms a

is asymmetric and that consumers are more sensitive to stock market falls is rejected by the data. The nature and significance of this alleged asymmetry has already proved to depend on data and model specification. This paper provides further evidence that linearity is indeed a reasonable working assumption in this context and, given that, consumption decisions are largely independent of stock market fluctuations.

single-threshold and a double-threshold piecewise model.

Appendix I –Cointegration tests.

These tests are implemented following Johansen (1988, 1991) and allowing for an intercept in both the cointegrating relationship and the VAR. All series except equity wealth display a linear trend, so that an intercept in the VAR is warranted. As for the cointegration vector, a non-trending relationship seems preferable in the light of both theoretical and statistical considerations: the theory implies a stationary equilibrium relationship (see section 2); if included, the trend is scarcely significant and it does not affect the other estimates. Ludvigson and Steindel (1999) adopt the same specification.

(1) non-durable consumption

VAR(1)

L-max		Trace		H_0
statistic	5% c.v.	statistic	5% c.v.	
41.56	27.07	64.27	47.21	0
12.92	20.97	22.71	29.68	1
9.76	14.07	9.79	15.41	2

VAR(2) *

L-max		Trace		H_0
Statistic	5% c.v.	statistic	5% c.v.	
39.56	27.07	59.23	47.21	0
13.72	20.97	19.67	29.68	1
5.85	14.07	5.94	15.41	2

VAR(3)

L-max		Trace		H_0
statistic	5% c.v.	statistic	5% c.v.	
33.29	27.07	51.18	47.21	0
13.40	20.97	17.88	29.68	1
4.47	14.07	4.48	15.41	2

VAR(4) **

L-max		Trace		H_0
Statistic	5% c.v.	statistic	5% c.v.	
27.94	27.07	46.83	47.21	0
13.85	20.97	18.89	29.68	1
5.02	14.07	5.04	15.41	2

* Optimal order according to Schwarz's information criterion
** Optimal order according to Akaike's information criterion

(2) total consumption

VAR(1)

L-max		Trace		H_0
statistic	5% c.v.	statistic	5% c.v.	
36.42	27.07	60.74	47.21	0
14.00	20.97	24.31	29.68	1
10.22	14.07	10.31	15.41	2

VAR(2) *

L-max		Trace		H_0
Statistic	5% c.v.	Statistic	5% c.v.	
22.44	27.07	48.57	47.21	0
19.20	20.97	26.13	29.68	1
6.77	14.07	6.93	15.41	2

VAR(3)

L-max		Trace		H_0
statistic	5% c.v.	statistic	5% c.v.	
22.91	27.07	41.82	47.21	0
13.16	20.97	18.90	29.68	1
5.70	14.07	5.74	15.41	2

VAR(4) **

L-max		Trace		H_0
Statistic	5% c.v.	statistic	5% c.v.	
21.76	27.07	39.19	47.21	0
11.18	20.97	17.42	29.68	1
6.22	14.07	6.23	15.41	2

* Optimal order according to Schwarz's information criterion
** Optimal order according to Akaike's information criterion

Appendix II – Local polynomial estimation.

Masry and Fan (1993) proved that the applicability of local estimation techniques to time series contexts depends on the stationarity and mixing properties of the data. Unit root tests provide strong evidence on the mean-stationarity of Δc^{nd}_t and Δwe_t. Since the latter is defined on the basis of a stock price index, in this case the stability of the variance is a concern as well. If the data is split in 5 sub-samples of roughly 7 years each, the standard deviations are respectively .030 ('67-'73), .034 ('74-'80), .036 ('81-'87), .025 ('88-'94), .037 ('95-'02). The statistics discussed by Levene (1960) and Brown and Forsythe (1974) allow to test the null hypothesis of constant variance without any assumption on the distribution of the random variable. With 5 sub-samples and 427 data points, the statistics have an $F_{(4,422)}$ distribution: they are respectively equal to 1.60 (p-value .17) and 1.33 (p-value .34); hence the null is not rejected. Finally, stationary ARMA sequences are strong mixing under mild regularity assumptions on the distribution of the error term (Athreya and Pantula, 1986). Hence, the central limit theorem established by Fan and Gijbels (1996) applies, and univariate local polynomial estimates are consistent and converge at the usual $\sqrt{T}$ rate. The *partial linear model* (PLM) in section 6 has the following form:

$$\Delta c^{nd}{}_t = \alpha + \mathrm{M}(L)\Delta c^{nd}{}_{t-1} + \mathrm{N}(L)\Delta y_{t-1} + \Pi(L)\Delta w^{ne}{}_{t-1} + \mathrm{T}(L) r_{t-1} + \sum_i f_i(\Delta w^{e}{}_{t-i}) + \varepsilon_t \qquad [11]$$

where M, N, Π, T are lag polynomials and the f_i are unknown real functions. It is possible to consider more general models where other explanatory variables besides w^e

enter non-linearly, or the f_i functions are not additively separable; such models, though, would incur all the problems connected to estimating non-parametrically in a multidimensional context. Given that the paper examines a specific proposition, namely that the link between equity prices and consumption is linear, the PLM above seems more appropriate.

The estimation is based on the iterative *Alternating Conditional Expectations (ACE)* algorithm introduced by Breiman and Friedman (1985). The only "extra" identification assumption with respect to the linear model is $E[f_i(\Delta w^e_{t-i})]=0 \ \forall i$, which avoids the introduction of free constants in the model. The f_i terms are estimated by a local linear smoother, applying the *LOWESS* method to robustify the results against potential outliers (Cleveland, 1979). The use of linear polynomials assumes that – in spite of being globally nonlinear – the f_i functions are sufficiently smooth to be considered linear locally, i.e. in a neighbourhood of each observation. The smoothing is based on the Epanechnikov kernel, whereas the *LOWESS* robustness weights are computed with a biweight kernel as originally suggested by Cleveland.

As in all kernel-based analyses, the choice of the bandwidth is fundamental. For every point x, the bandwidth determines the size of the neighbourhood N_x on which the estimation of $f(x)$ is based; in choosing it, a well known trade-off between bias and variance has to be faced: a small bandwidth implies a low bias and a large sample variability. Given the extremely uneven distribution of Δwe_t, I use an adaptive "nearest neighbour" bandwidth. The bandwidth is adaptive in the sense that it is inversely proportional to the local density of the regressor: a large (small) bandwidth is used in regions where the data are sparse (dense), so that a constant number of observations k is taken into account in each estimation. Hence, the degree of smoothing is regulated by the parameter k. The asymptotically optimal $k^*=350$ is derived by minimising the *cross-validation* function $CV(k)$, which is a consistent estimator of the mean integrated square error *(MISE)*, namely the mean square error integrated over the domain of the regressor Δwe_{t-1}. Throughout the ACE iterations, a smaller bandwidth $(k'=250)$ is deliberately used in order to improve the accuracy of the linear coefficients estimates (Fan and Gjibels, 1996). After achieving convergence, an extra step is implemented using k^* to generate the estimates for the f_i functions. Sensitivity analysis confirms that the behaviour of the estimates is a regular function of k and does not depend on the (arbitrary) choice of k'.

The estimation is implemented in MatLab with author's codes. An extensive discussion of local polynomial modelling can be found in Fan and Gjibels (1996), together with more detailed references.

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Table 1 – VECM(3) representation

	Δc^{nd}_t	Δy_t	Δw^e_t	Δw^{ne}_t
ec_{t-1}	-.00 *(-1.37)*	.02 *(.92)*	.41 *(4.31)*	.00 *(.70)*
Δc^{nd}_{t-1}	-.42*	.17	.26	-.04
Δc^{nd}_{t-2}	-.13*	.20	-.27	.03
Δc^{nd}_{t-3}	.14*	.44*	-.37	.07*
Δy_{t-1}	.02	-.32*	.17	-.03
Δy_{t-2}	.05	-.27*	.03	-.05*
Δy_{t-3}	.02	-.23*	-.27	-.02
Δw^e_{t-1}	.00	.00	-.15*	-.00
Δw^e_{t-2}	.00	.00	-.15*	.00
Δw^e_{t-3}	.00	.01	.00	.01*
Δw^{ne}_{t-1}	.17*	.22	1.10	.97*
Δw^{ne}_{t-2}	.00	-.01	2.19*	-.13*
Δw^{ne}_{t-3}	.09	.03	1.14	-.18*
R^2_{adj}	.21	.15	.21	.64
SE	.003	.007	.029	.002

c^{nd}: non-durable consumption; y: labour income ; w^e: equity wealth; w^{ne}: non-equity wealth;
ec: cointegration residual from equation [9], namely $c^{nd} + 2.1 - .93y - .04w^e - .21w^{ne}$.
All variables are logarithms of real per-capita series. Asterisks denote significance at the 5% level.

Table 2 – Linear model for Δc^{nd}

C	dummies		Δc^{nd}_{t-i}		Δy_{t-i}		Δw^{ne}_{t-i}		r^{j}_{t-i}		Δw^{e}_{t-i}	
.003	a	-.003	i=1	-.491	i=2	.044	i=1	.096	r^1_{t-2}	-.004	i=10	.015
(.000)		*(.001)*		*(.046)*		*(.021)*		*(.046)*		*(.001)*		*(.005)*
	b	-.003	i=2	-.191	i=8	.054	i=3	.086	r^1_{t-6}	.002		
		(.001)		*(.051)*		*(.021)*		*(.048)*		*(.001)*		
	c	-.003	i=3	.118			i=6	.078	r^1_{t-12}	.004		
		(.001)		*(.046)*				*(.043)*		*(.002)*		
	d	-.003							r^2_{t-10}	.004		
		(.001)								*(.002)*		
	e	-.008							r^2_{t-12}	-.007		
		(.003)								*(.002)*		
									r^{sp}_{t-8}	.005		
										(.002)		

sum of squared residuals: .004; standard errors in brackets.

a : 1973/8-1974/3	c^{nd} : non-durable consumption	r^1 : 1 month certificate of deposit rate
b : 1979/12-1980/4	y : labor income	r^2 : 10 years treasury bill rate
c : 1990/7-1991/3	w^{ne}: non-equity wealth	r^{sp}: r^2-r^1
d : 1992/4-1992/8	w^e : equity wealth	
e : 2001/9		

Table 3 – Non-linear model for Δc^{nd}

C	*dummies*		Δc^{nd}_{t-i}		Δy_{t-i}		Δw^{ne}_{t-i}		r^{j}_{t-i}		$f(\Delta w^{e}_{t-10})$
.003	a	-.003	i=1	-.496	i=2	.046	i=1	.111	r^1_{t-2}	-.004	*[see figures 6 and 7]*
(.000)		*(.001)*		*(.047)*		*(.022)*		*(.047)*		*(.001)*	
	b	-.003	i=2	-.197	i=8	.057	i=3	.107	r^1_{t-6}	.003	
		(.002)		*(.052)*		*(.021)*		*(.048)*		*(.001)*	
	c	-.003	i=3	.100			i=6	.083	r^1_{t-12}	.004	
		(.001)		*(.047)*				*(.044)*		*(.002)*	
	d	-.004							r^2_{t-10}	.005	
		(.002)								*(.002)*	
	e	-.010							r^2_{t-12}	-.008	
		(.003)								*(.002)*	
									r^{sp}_{t-8}	.005	
										(.002)	

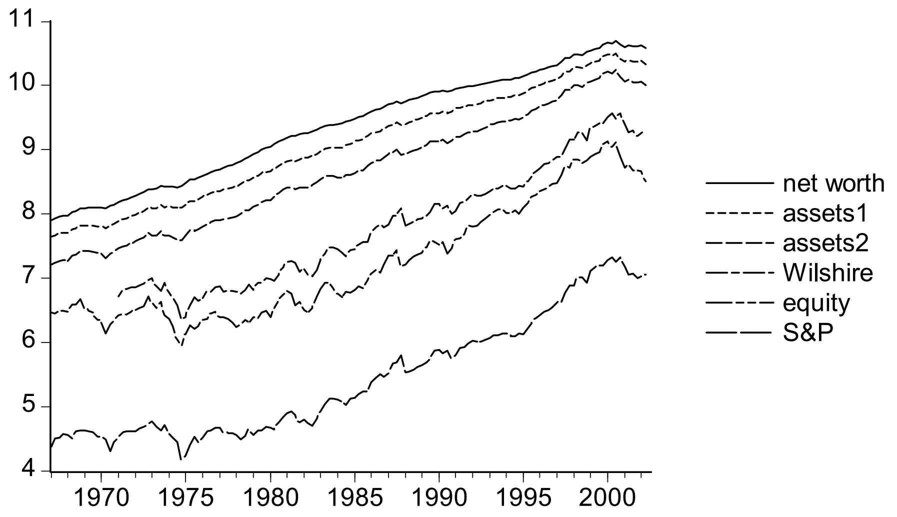
sum of squared residuals: .004; standard errors in brackets.

a : 1973/8-1974/3
b : 1979/12-1980/4
c : 1990/7-1991/3
d : 1992/4-1992/8
e : 2001/9

c^{nd} : non-durable consumption
y : labor income
w^{ne}: non-equity wealth
w^{e} : equity wealth

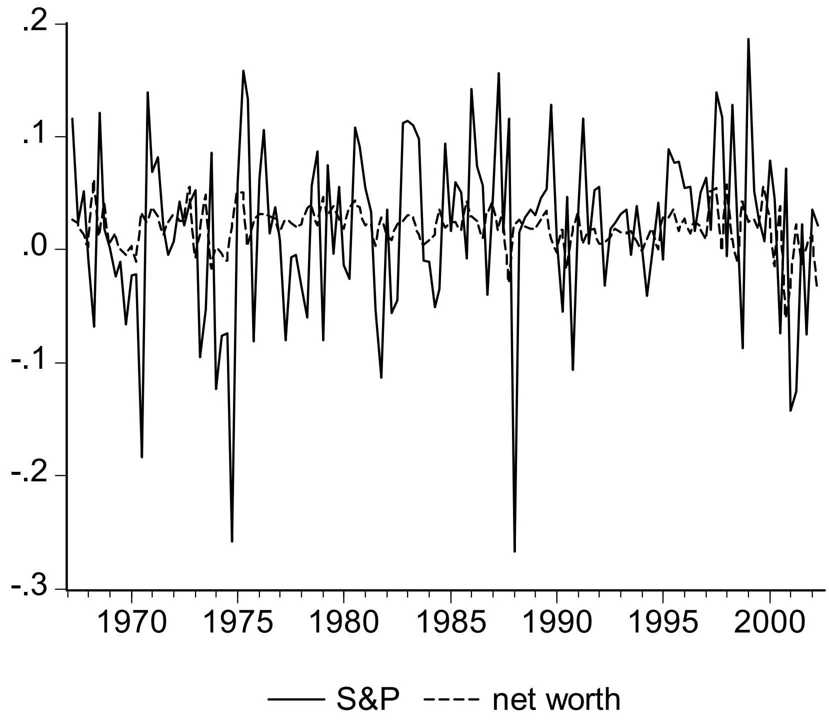
r^1 : 1 month certificate of deposit rate
r^2 : 10 years treasury bill rate
r^{sp}: r^2-r^1

Figure 1 – Wealth and equity prices



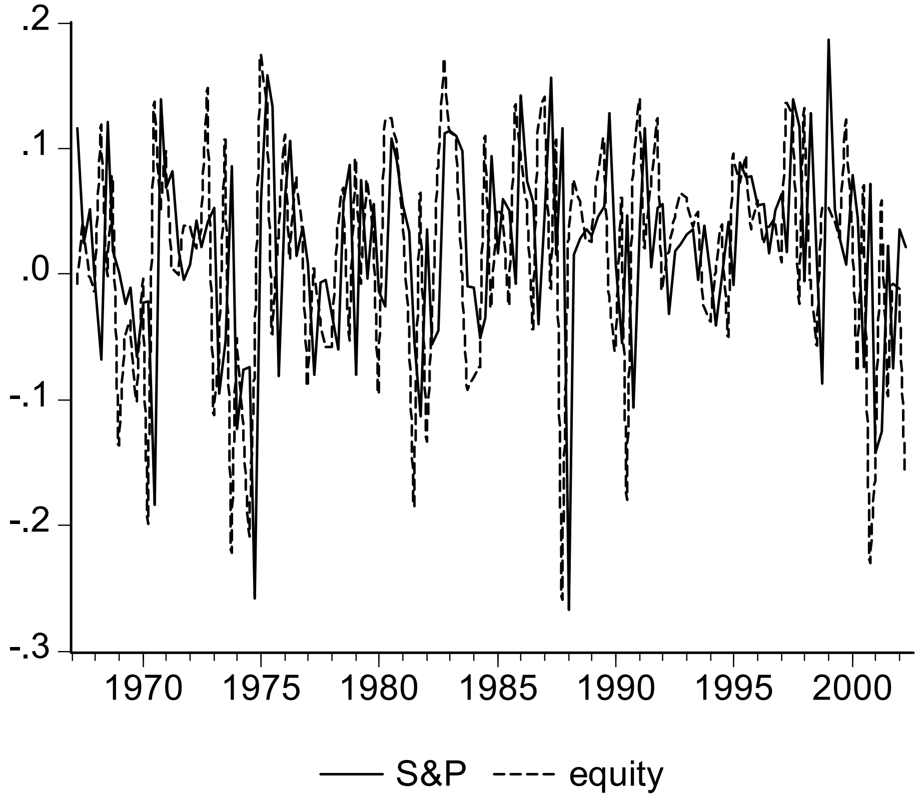
Net Worth: BEA net worth. *Assets1*: BEA financial assets. *Assets2*: corporate equities, mutual fund shares, pension fund reserves, investment in bank personal trusts, equity in non-corporate business. *Wilshire*: Wilshire 5000 equity price index. *Equity*: corporate equities. *S&P*: Standard&Poors 500 equity price index. Logarithms of seasonally-adjusted quarterly series. Source: Federal Reserve and Datastream.

Figure 2 – Growth rates of net worth and equity prices



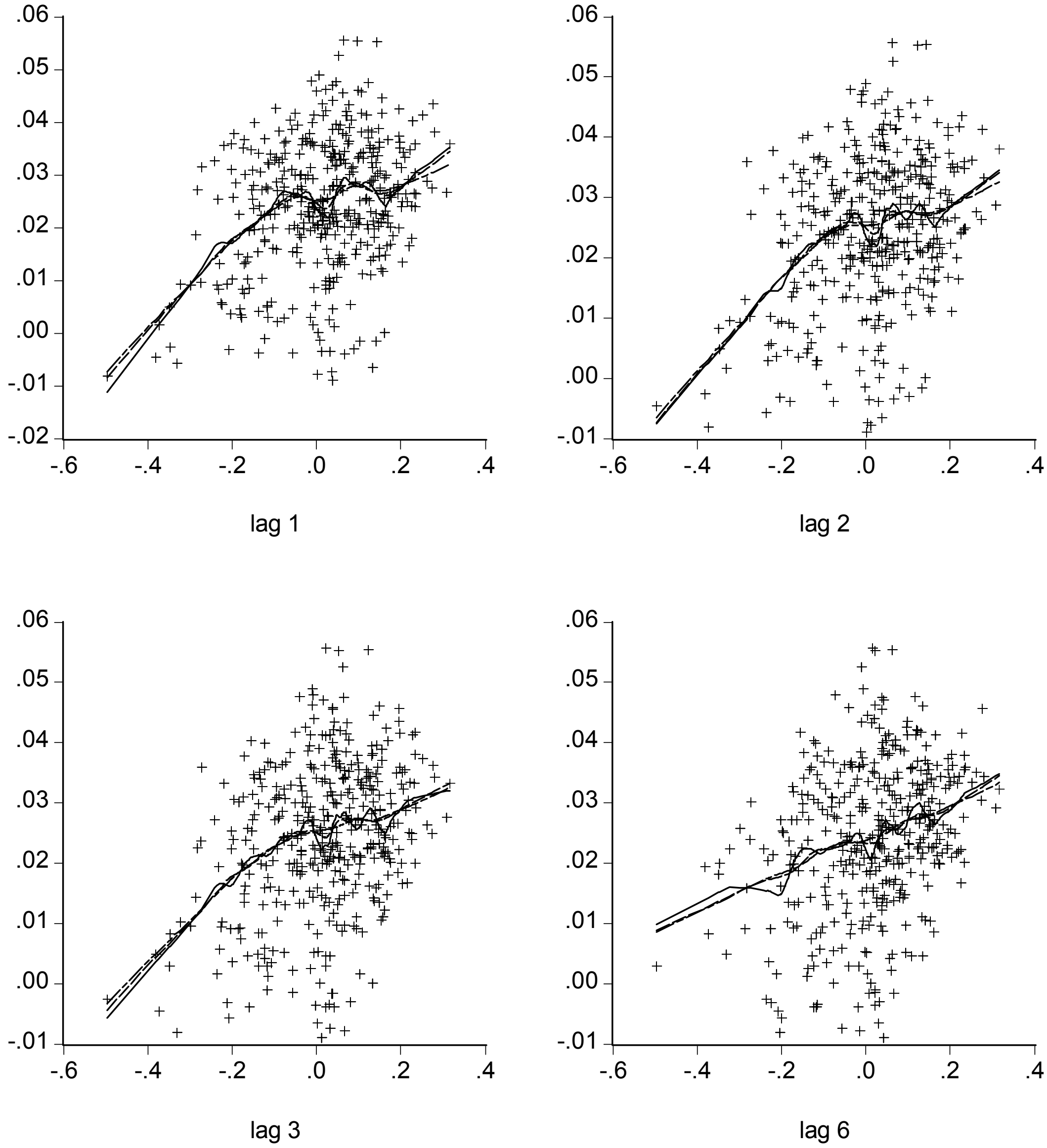
S&P: Standard&Poors 500 equity price index, quarterly growth rate. *Net worth*: BEA net worth, quarterly growth rate. The growth rates are obtained as first differences of logarithms of seasonally-adjusted quarterly series. Source: Federal Reserve and Datastream.

Figure 3 – Growth rates of corporate equity holdings and equity prices



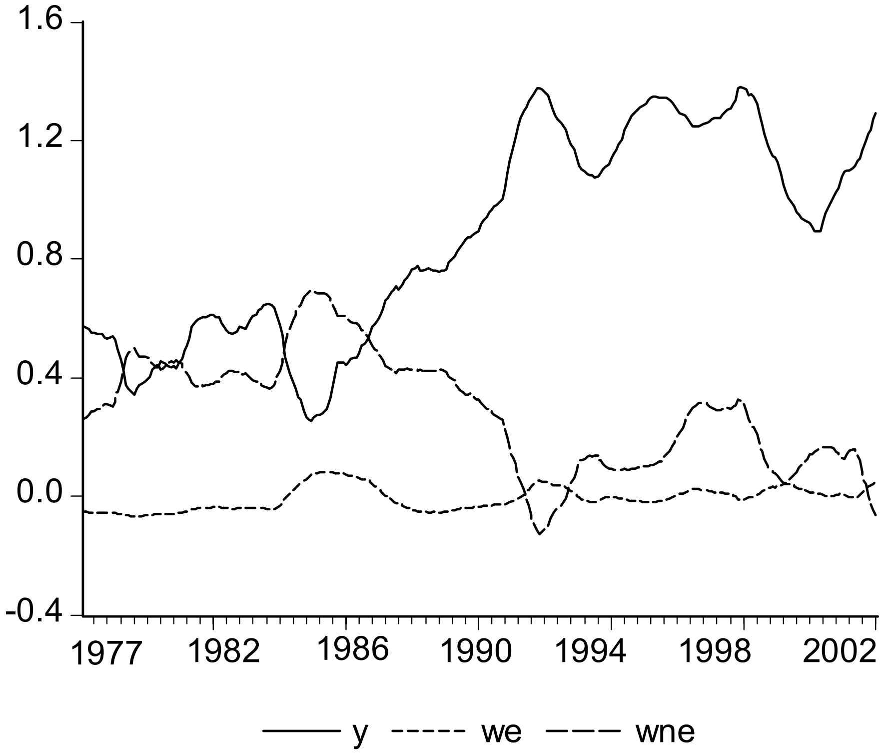
S&P: Standard&Poors 500 equity price index, quarterly growth rate. *Equity*: BEA corporate equity, quarterly growth rate. The growth rates are obtained as first differences of logarithms of seasonally-adjusted quarterly series. Source: Federal Reserve and Datastream.

Figure 4 – Consumption growth as a function of past equity wealth growth



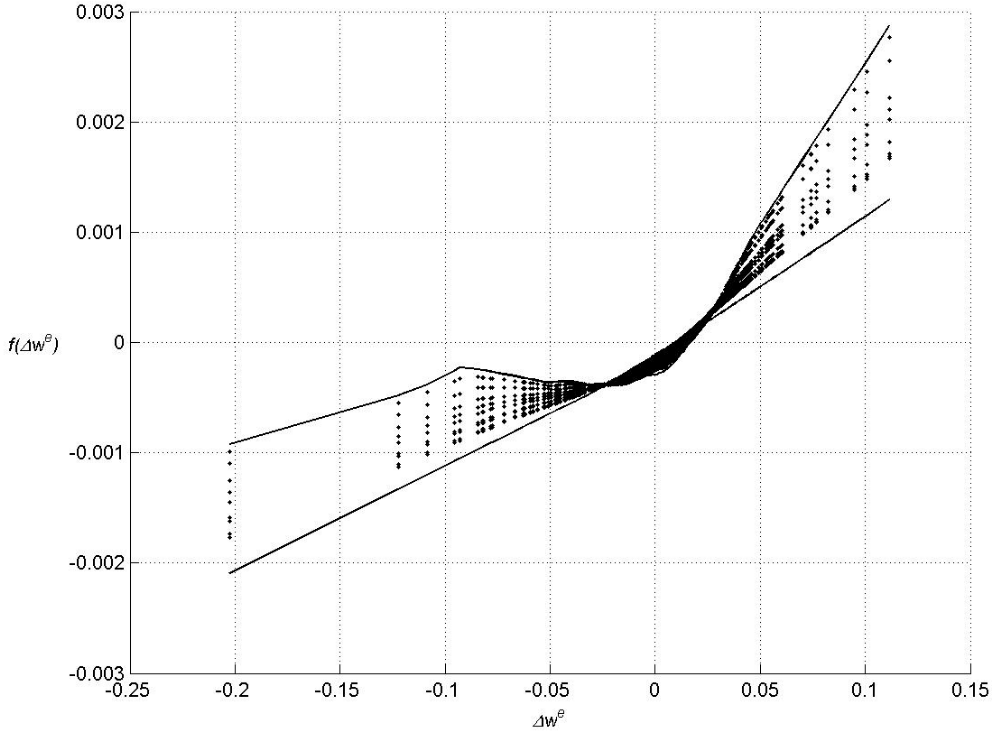
Estimates of the function f in equation [7]: $g^{c}_{t} = f(g^{we}_{t-i}) + \varepsilon_t$, where g^{we}_{t-i} is the annual growth rate of equity wealth $(= \Delta_{12} we_{t-i})$ and g^{c}_{t} is the annual growth rate of non-durable consumption $(= \Delta_{12} c^{nd}_{t})$. Each plot shows the LOESS estimate of f for a different lag i=1,2,3,6, based on local linear polynomials, tricube kernel and bandwidths of .15 (continuous line), .30 (short dash) and .45 (long dash).

Figure 5 – The pattern of long-run consumption elasticities



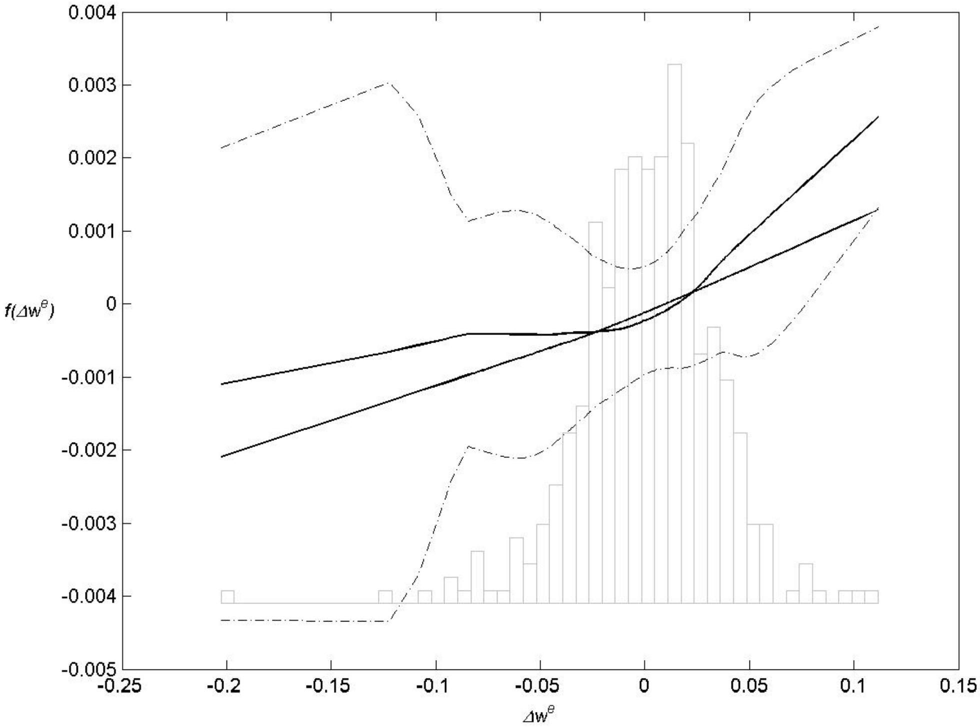
Rolling estimates of the elasticity of non-durable consumption with respect to labour income *(y)*, equity wealth *(we)* and non-equity wealth *(wne)*. Starting from January 1977, the time-t parameters are obtained by estimating the cointegrating relationship between consumption, income and wealth (equation [9]) using the previous 10 years of data.

Figure 6 – $f(\Delta w^{e}_{t-10})$ at different bandwidths



Horizontal axis: Δw^{e}_{t-10}. Vertical axis: local polynomial estimates of $f(\Delta w^{e}_{t-10})$ based on alternative bandwidths. The straight line represents the extreme case where the bandwidth covers the whole sample and the local polynomial estimate coincides with the OLS regression line.

Figure 7 – Confidence band for *$f(\Delta w^{e}_{t-10})$*



Horizontal axis: Δw^{e}_{t-10}. Vertical axis: local polynomial estimate of $f(\Delta w^{e}_{t-10})$ based on an optimal bandwidth selected by cross-validation, with 95% point-wise confidence intervals. The background histogram is the empirical distribution of Δw^{e}_{t-10}.