The U.S. consumption-wealth ratio and predictability of national stock returns: Evidence for the G7

Thomas Nitschka*
Department of Economics, University of Dortmund
D-44221 Dortmund, Germany
E-mail: T.Nitschka@wiso.uni-dortmund.de
Phone: ++49-(0)231-755-3266
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Abstract

This paper contributes to the growing body of empirical literature on long-term predictability of expected stock returns. It is followed Lettau and Ludvigson (2001 JoF, 2004 AER) who discovered that a trivariate empirical approximation of the logarithmic consumption wealth ratio cointegrates in U.S. data. They found that the cointegration residual reveals predictive power for excess and real returns on U.S. stock market indizes.

Evidence presented in this paper suggests that a four-variable approximation of the U.S. consumption-wealth ratio cointegrates and

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embodies information about changes of foreign stock market wealth. It is shown that the respective cointegration residual is a predictor of market capitalization changes of foreign stock indizes at 14-20 quarters horizon. This finding is robust irrespective if market capitalizations are denoted in current U.S. dollars or local currency.

1 Introduction

Long-term predictability of asset returns is well documented in a growing body of empirical literature[1]. This paper contributes to that literature by employing the method proposed by Lettau and Ludvigson (2001 JoF, 2004 AER) to address the question if the U.S. consumption-wealth ratio is not only informative about the future path of U.S. but also foreign stock indizes. Lettau and Ludvigson (henceforth L&L) discovered that a trivariate empirical approximation of the consumption wealth ratio cointegrates in U.S. data.[2] They found that the cointegration residual reveals predictive power for excess and real returns on U.S. stock market indizes because transitory deviations from the common trend are mainly induced by the stock market component of U.S. households´ asset wealth. As U.S. households either directly or indirectly hold foreign equity, the U.S. consumption wealth ratio should embody information about movements of foreign stock markets.

Evidence presented in this paper suggests that a four-variable logarithmic approximation of the U.S. consumption-wealth ratio contains information about changes of foreign equity holdings in U.S. stock market wealth. It is shown that the respective cointegration residual is a powerful predictor of market capitalization changes of foreign stock indizes irrespective if these are expressed in current U.S. dollars or local currency. Bilateral U.S. dollar exchange rates seem to play a negligible role in this context. Thus fluctuations of the U.S. consumption-wealth ratio do not only mirror movements of U.S. but also foreign stock markets.

[1] see Fama and French (1988a,1988b), Poterba and Summers (1988), Campbell and Shiller (1988) for U.S. data and Campbell and Hamao (1991) as well as Richards (1995) for international data and Santos and Veronesi (2004), Piazzesi et al. (2004) as well as Lettau and Ludvigson (2001,2004) for macroeconomically founded predictions of (excess) stock returns in U.S. data.

[2] Fernandez-Corugedo et al. (2003) directly employ the method of Lettau and Ludvigson which corroborates their findings in U.K. data.

The remainder of this paper is organized as follows. A simple manipulation of the theoretical framework used by L&L is introduced in section two. Section three provides details on forecast regressions of market capitalization changes expressed in current U.S. dollars and national currency for G7 MSCI stock market indizes. Furthermore, it is analysed if changes of the bilateral U.S. dollar spot exchange rates are responsible for the forecast regression results for market capitalization changes denoted in current U.S. dollars. Section four concludes. A detailed description of the data employed in this paper is given in the appendix.

2 The Consumption-Wealth Ratio

The forecast ability of the logarithmic consumption-wealth ratio for returns on broad U.S. stock indizes seems to be predominantly driven by temporary deviations of U.S. households´ stock market wealth from the common trend among consumption, asset wealth and labour income (L&L 2004). Moreover, U.S. households do not only invest in U.S. corporate equity but hold at least to a small extent either directly or indirectly foreign equity. Investment in foreign equity is very small relative to total wealth but inevitably raises the question if the U.S. consumption-wealth ratio embodies information about fluctuations of foreign stock markets and hence on returns on foreign stock indizes. A simple manipulation of the theoretical framework of L&L which allows to explicitly deal with this issue is presented next.

2.1 Model Setup

This subsection presents a simple manipulation of the approach employed by L&L who follow Campbell and Mankiw (1989) and regard a representative agent economy in which all wealth is traded. W_t denotes aggregate wealth (human capital plus asset wealth) in period t. C_t denotes consumption and $r_{w,t+1}$ the net return on aggregate wealth. Thus, the budget constraint of the representative household can be written as

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

Dividing by W_t and taking logs of (1) gives (2), where $r_{w,t+1}$ now denotes the logarithmic approximation of the net return on aggregate wealth. In the following lower-case letters denote logarithms.

$$w_{t+1} - w_t = r_{w,t+1} + \log(1 - \frac{C_t}{W_t}) \tag{2}$$

$$w_{t+1} - w_t = r_{w,t+1} + \log(1 - \exp(c_t - w_t)) \tag{3}$$

Assuming that the log consumption-wealth ratio is stationary, a long run mean of $c_t - w_t$, $\overline{c - w}$, exists. Then the last term of the right-hand side of (3) can be approximated via a Taylor expansion around the steady state consumption-wealth ratio. Rearranging, summarizing all constant elements and denoting them by κ as well as substituting $1 - \exp(\overline{c - w})$ for ρ_w gives

$$\log(1 - \exp(c_t - w_t)) \approx \kappa + (1 - \frac{1}{\rho_w})(c_t - w_t) \tag{4}$$

Plugging (4) into (3) and additionally exploiting that $w_{t+1} - w_t = \Delta w_{t+1}$ one obtains

$$\Delta w_{t+1} \approx \kappa + r_{w,t+1} + (1 - \frac{1}{\rho_w})(c_t - w_t) \tag{5}$$

It is also possible to write Δw_{t+1} tautologically in terms of the consumption growth rate and changes of the consumption-wealth ratio

$$\Delta w_{t+1} = \Delta c_{t+1} + (c_t - w_t) - (c_{t+1} - w_{t+1})$$

Substitution in (5) gives

$$c_t - w_t = \rho_w(r_{w,t+1} - \Delta c_{t+1}) + \rho_w(c_{t+1} - w_{t+1}) + \rho_w\kappa \tag{6}$$

Solving forward to the infinite horizon, neglecting constant terms, taking expectations and imposing a transversality condition, $\lim_{i\to\infty} \rho_w^i(c_{t+i+1} - w_{t+i+1}) = 0$, leads to the following expression for the log consumption-wealth ratio

$$c_t - w_t = E_t \sum_{i=1}^{\infty} \rho_w^i (r_{w,t+i} - \Delta c_{t+i}) \tag{7}$$

According to this equation fluctuations of the log consumption-wealth ratio either display variation of expected returns on aggregate wealth or expected changes of consumption. However, (7) cannot be employed for empirical purposes because one part of aggregate wealth, human capital, is unobservable.

This issue can be solved by assuming that aggregate labour income is the dividend paid from human capital and represents the non-stationary component of human capital. The gross return on human capital is defined as

$$1 + r_{h,t+1} = \frac{H_{t+1} + Y_{t+1}}{H_t} \tag{8}$$

where H_t denotes the level of human capital and Y_t the level of labour income at time t. Solving (8) for H_t gives

$$H_t = \frac{H_{t+1} + Y_{t+1}}{1 + r_{h,t+1}} \tag{9}$$

Expanding (9) to the infinite horizon and taking expectations leads to

$$H_t = E_t[\sum_{j=1}^{\infty} \prod_{i=1}^{j} (1 + r_{h,t+i})^{-i} Y_{t+j} \tag{10}$$

which says that human capital is the present discounted value of expected labour income. If one returns to the one-period scenario and takes logarithms of (8) it is possible to employ the Campbell-Shiller return decomposition (Campbell and Shiller (1988)) under the assumption that labour income as well as human capital are integrated of order one, I(1), and the ratio of log labour income with log human capital is stationary, which gives

$$\begin{aligned} r_{h,t+1} &= \Delta h_{t+1} + K + (1 - \rho_h)(y_{t+1} - h_{t+1}) \\ &= -h_t + \rho_h h_{t+1} + (1 - \rho_h) y_{t+1} + K \end{aligned} \tag{11}$$

with K representing constant elements that are obtained in the course of the return decomposition and $\rho_h \equiv \frac{1}{1+\exp(\overline{y-h})}$, where $\overline{y-h}$ denotes the long-run mean of the log labour income-log human capital ratio. Solving (11) for h_t and extending to the infinite horizon as well as subtracting y_t on both sides of the equation leads to

$$
\begin{aligned}
h_t - y_t &= \frac{K}{1-\rho_h} + E_t \sum_{j=1}^{\infty} \rho_h^j (\Delta y_{t+j} - r_{h,t+j}) && (12) \\
h_t &= y_t + \frac{K}{1-\rho_h} + E_t \sum_{j=1}^{\infty} \rho_h^j (\Delta y_{t+j} - r_{h,t+j}) && (13)
\end{aligned}
$$

if expectations are taken on both sides of the equation. Exploiting the assumption that labour income is integrated of order one equation (13) gives an expression of human capital in terms of a constant, $\kappa = \frac{K}{1-\rho_h}$, log aggregate labour income, y_t, and a covariance stationary term,

$$
z_t = E_t \sum_{j=1}^{\infty} \rho_h^j (\Delta y_{t+j} - r_{h,t+j})
$$

such that

$$
h_t = \kappa + y_t + z_t
$$

This equation is employed to express the unobservable variable h_t in terms of observable variables. In addition, aggregate wealth can be decomposed into its components asset and human wealth

$$
W_t = A_t + H_t
$$

with A_t representing asset wealth and H_t human wealth. Then log aggregate wealth can be approximated around the steady state by

$$
w_t \approx v a_t + (1-v) h_t \tag{14}
$$

with v interpretable as steady state share of asset wealth in aggregate wealth. Furthermore, employing the same technique log asset wealth is expressed as function of its components, here $A_t = FW_t + DAW_t$, where FW_t represents foreign equity holdings and DAW_t denotes that part of aggregate wealth that is not foreign equity or human capital which is referred to as domestic asset wealth. Thus the logarithmic approximation of asset wealth obeys

$$
a_t \approx \lambda f w_t + (1-\lambda) da w_t \tag{15}
$$

with λ representing the steady state share of foreign equity holdings in asset wealth. Combining these two results gives

$$
w_t \approx \theta f w_t + \phi da w_t + (1-\theta-\phi) h_t \tag{16}
$$

with $\theta = v\lambda = \frac{\overline{FW}}{\overline{W}}$ the steady state share of foreign equity in total wealth, $\phi = v(1-\lambda) = \frac{\overline{DAW}}{\overline{W}}$ the steady state share of domestic asset wealth in aggregate wealth.

Additionally, one could decompose the gross return on aggregate wealth into the returns on its components, here first into the components asset wealth and human capital which gives

$$1 + r_{w,t} = v_t(1 + r_{a,t}) + (1 - v_t)(1 + r_{h,t}) \tag{17}$$

Campbell (1996) proved that taking logarithms of (17) reduces the equation to

$$r_{w,t} = v r_{a,t} + (1 - v) r_{h,t} \tag{18}$$

Again the same technique can be employed to receive a logarithmic approximation of the return on asset wealth in terms of the returns on its components foreign equity and domestic asset wealth that can be combined with equation (18)

which leads to

$$r_{w,t} = \theta r_{fw,t} + \phi r_{daw,t} + (1 - \theta - \phi) r_{h,t} \tag{19}$$

Plugging (19) and (16) into (7) gives

$$\begin{aligned} & c_t - \theta f w_t - \phi daw_t - (1 - \theta - \phi) h_t \qquad (20) \\ = \quad & E_t\{\sum_{i=1}^{\infty} \rho_w^i [(\theta r_{fw,t+i} + \phi r_{daw,t+i} + (1 - \varphi - \phi) r_{h,t+i}) - \Delta c_{t+i}]\} \end{aligned}$$

The unobserved variable h_t still occurs on the left-hand side but can be replaced by the expression for h_t derived above assuming that $\rho_w^i = \rho_h^i$.

$$\begin{aligned} & c_t - \varphi f w_t - \phi daw_t - (1 - \varphi - \phi) y_t \qquad (21) \\ = \quad & E_t\{\sum_{i=1}^{\infty} \rho_w^i [(\varphi r_{fw,t+i} + \phi r_{daw,t+i} + (1 - \varphi - \phi)\Delta y_{t+i}) - \Delta c_{t+i}] + (1 - \varphi - \phi) z_{t+i}\} \end{aligned}$$

According to (21), c_t, log consumption, fw_t, log foreign equity holdings, daw_t, log domestic asset wealth and y_t, log labour income, are cointegrated as all the variables on the right-hand side should be stationary if the variables on the left-hand side are integrated of order one which is tested below. Hence, time variation of the consumption-wealth ratio, i.e. a deviation from the common long-term trend, should either mirror changes of (returns on) foreign equity holdings, changes of domestic asset wealth, changes of labour income or consumption growth, or a combination of these.

2.2 Empirical evidence: Cointegration and error correction

This section assesses the question if the four-variable proxy for the log consumption-wealth ratio proposed above cointegrates. All variables employed are real, per capita, expressed in billions of chain-weighted 2000 dollars for the sample period from second quarter 1952 to second quarter 2004. It is followed Blinder and Deaton (1985) who suggest to proxy total consumption as constant multiple of non-durables and services consumption expenditure excluding clothing and footwear. Labour income is proxied as proposed by L&L (2001,2004). U.S. households´ foreign equity holdings are determined as explained in detail in the appendix. Domestic asset wealth is calculated as household net worth less foreign equity holdings. Augmented Dickey-Fuller test results provide evidence that each variable employed in this analysis contains a unit root. Furthermore, it cannot be rejected that first differences of these variables are stationary.[3] Hence, all the variables are integrated of order one, which suggests that the approximation of the consumption-wealth ratio derived above should cointegrate. Results of the Johansen cointegration test are displayed in table 1. The table shows that Akaike(AIC) and Schwartz (SIC) information criteria suggest an appropriate lag length of one for the vector autoregressive representation (VAR) of the four variables under consideration. Table 1 presents critical values for Trace and L-max test as well as the test statistics for both tests. Formally, one cannot reject the null of no cointegration for the relation between non-durables and services consumption expenditure excluding clothing and footwear, foreign equity holdings, domestic asset wealth and labour income at 90% confidence level. However, theory as well as unit root tests suggest the presence of cointegration.[4] More-

[3] results available upon request

[4] Hoffmann and Mc Donald (2003) show that the existence of a cointegration relationship cannot be only grounded on statistical terms but should incorporate economic theory.

over, estimates of the cointegration vector are highly plausible with respect to theory, which is addressed below. That is why i believe that cointegration among non-durables and services consumption[5], foreign equity holdings, domestic asset wealth and labour income exists which is assumed to be the case throughout the remainder of the paper.

One reason to believe that cointegration is present are the economically meaningful estimates of the cointegration vector calculated below.

As emphasized by Stock (1987) the OLS estimates of cointegrated variables converge to their true value with the sample size rather than with the square root of the sample size. Thus these estimates are "superconsistent" and simple OLS provides consistent point estimates. However, the error terms of the individual time-series variables could be correlated with each other. Hence the OLS estimates are consistent but could be substantially biased away from the true values because of the above mentioned second-order bias. That is why i follow Stock and Watson (1993) who propose a dynamic least squares technique to overcome this obstacle, which is achieved by adding leads and lags of the first differences of foreign equity holdings, domestic asset wealth and labour income. Hence the estimate equation takes the following form

$$\begin{aligned} c_t &= \alpha + \beta_{fw} fw_t + \beta_{daw} daw_t + \beta_y y_t \qquad (22) \\ &\quad + \sum_{i=-k}^{k} b_{fw,i} \Delta fw_{t-i} + \sum_{i=-k}^{k} b_{dw,i} \Delta daw_{t-i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t-i} + \varepsilon_t \end{aligned}$$

The estimation of the cointegration coefficients gives the following results if the coefficient on non-durable consumption is normalized to unity with t-statistics in parentheses. The coefficients of the differences in lead or lag are omitted.[6]

$$\widehat{\beta} = [1 \; - \underset{(2.8954)}{0.0106} fw_t \; - \underset{(8.9507)}{0.3409} daw_t \; - \underset{(25.3801)}{0.7331} y_t]' \qquad (23)$$

[5] Rudd and Whelan (2002) criticize the use of non-durable consumption expenditures because the budget constraint refers to total personal consumption. In addition, they provide arguments that log total consumption and log non-durable and services consumption are not linearily linked over time. However, L&L argue that durable goods represent a stock of goods and hence are better described as wealth which is the view that is followed in this paper.

[6] The estimates do not vary much from one to seven leads and lags. Here six leads and lags are employed.

At first glance the estimated cointegration coefficients do not seem to be economically meaningful as they sum to a number bigger than one. However, the reason for this is that only a share of total consumption is used in the estimation. It is assumed that total personal consumption is a constant multiple of non-durables and services consumption, i.e. total personal consumption less consumption of durable goods on the left hand side. However, durable goods are included in the asset wealth proxy on the right hand side such that the estimates should sum to a number bigger than one. Here they sum to 1.0846. The average share of durable goods in aggregate wealth is around 8%. Hence, the estimated cointegration vector is economically plausible.

Furthermore, taking the reasoning from above into account, the cointegration coefficient estimates mirror that the steady state share of labour income in wealth is roughly 0.7 and that of asset wealth approximately 0.3. This corroborates the results of L&L (2001,2004). Moreover, assuming that aggregate wealth represents output which is assumed to be governed by a Cobb-Douglas production function the cointegration coefficients could be interpreted as reflecting the shares of capital and labour in output which are stable over time. Translating the cointegration coefficients into shares of capital and labour, the share of capital would be 0.3. A number close to values employed in the real business cycle literature.[7] The point estimate of foreign equity holdings cointegration coefficient seems to be reasonable as well. The share of foreign equity in U.S. household net worth considerably increased since the 1990s but was virtually zero in the 1950s and 60s. Therefore these estimates make sense economically. As already emphasized above, based on these results i assume the presence of cointegration between the four variables under consideration throughout the remainder of the paper.

However, in order to answer the question whether deviations from the cointegration trend reflect transitory, predictable, movements in foreign equity holdings, domestic asset wealth, consumption or labour income the fundamental insight is employed that for every cointegration relation an error-correction representation exists (Engle and Granger (1987)).

The vector error correction representation (VECM) of $x_t = (c_t, fw_t, daw_t, y_t)'$ is

$$\Delta x_t = \mu + \alpha\widehat{\beta}' x_{t-1} + \Gamma(L)\Delta x_{t-1} + \varepsilon_t$$

where $\Delta x_t = (\Delta c_t, \Delta fw_t, \Delta daw_t, \Delta y_t)'$ is the vector of first differences

[7] see e.g Kydland, F.; Prescott, E. (1982)

and Δx_{t-1} the vector of lagged differences respectively, μ is a(4x1) vector of constants, $\alpha \equiv (\alpha_c, \alpha_{fw}, \alpha_{dw}, \alpha_y)'$ is the vector of adjustment coefficients which reflect what variable is responsible for the error correction. $\Gamma(L)$ denotes the lag operator matrix and $\widehat{\beta} \equiv (1, -\widehat{\beta}_{fw}, -\widehat{\beta}_{daw}, -\widehat{\beta}_y)'$ represents the vector of the above estimated cointegration coefficients. Hats indicate estimated variables and ε_t represents the vector of shocks in the cointegration relation.

The term $\widehat{\beta}' x_{t-1}$ gives the cointegration residual, α is the adjustment vector that displays what variables adjust a deviation from the common trend. This is one conclusion that can be drawn from the Granger Representation Theorem: If x_t is cointegrated, at least one of the adjustment coefficients $\alpha_c, \alpha_{fw}, a_{dw}$ or α_y must be nonzero in the error-correction representation.

All coefficients are estimated by OLS applying a lag length of one which is suggested by Akaike and Schwartz information criteria. As only the adjustment coefficients are of importance in this context other coefficient estimates of the VECM are omitted. The t-statistics of the adjustment coefficient estimates are reported in parentheses.

$$\alpha \equiv (\underset{(-0.8541}{-0.01183}, \underset{(4.3329)}{1.2986}, \underset{(3.7045)}{0.2252}, \underset{(-0.1429)}{-0.0043})'$$

Apparently both asset wealth components are responsible for a restoration of the common trend which could be expected regarding the results of L&L (2001,2004). Domestic asset wealth also adjusts to the common cointegration trend which is persumably driven by the domestic stock market wealth component. Furthermore, the foreign equity holdings adjustment coefficient is not only relatively high but also statistically significant. Hence the conclusion can be drawn that the U.S. consumption-wealth ratio embodies information on temporary changes of U.S. households´ foreign equity holdings. Thus the cointegration residual should serve as predictor of expected changes of the rest-of-the world equity position of U.S. households.

3 Forecasting power of the cointegration residual

In this section evidence for the predictive power of the cointegration residual, $\widehat{\beta}' x_{t-1}$, for movements of foreign stock markets is presented.

The results of different forecast regressions are summarized in tables 2, 3 and 4.[8] The coefficients of the regressor, $\widehat{\beta}' x_{t-1}$, with Newey-West corrected t-statistics as well as the adjusted R^2 are reported. The forecast horizon, h, is in quarters.

Before describing the evidence it may be useful to provide some economic intuition of what should be reflected in the regression outcomes. According to equation (7), a temporarily high consumption-wealth ratio should either mirror high expected returns on aggregate wealth or low expected consumption growth. L&L could show that asset wealth, in particular stock market wealth, is responsible for transitory deviations from the common trend among consumption and wealth. Consumption growth is not predictable. Hence, a high consumption-wealth ratio, i.e. a positive deviation from the cointegration trend should be associated with the expectation of high future returns on wealth, especially on stock market wealth. If changes of stock index market capitalizations are interpreted as returns on the respective index under consideration, then the reasoning from above should be reflected in positive regressor estimates in the forecast regressions.

Panel A of table 2 presents evidence from regressions of changes of U.S. households´ foreign equity holdings, ΔRoW, and of changes of the market capitalization of a value-weighted rest-of-the world index, $\Delta RoW - Index$, constructed with data on the MSCI stock market capitalization of the United Kingdom, France, Italy, Canada, Japan and Germany in current U.S. dollars.

The results displayed in the first column of Panel A allow to conclude that the cointegration residual is a powerful predictor of changes of the rest-of-the-world equity holdings of U.S. households. This sample spans the second quarter of 1952 to the second quarter of 2004. The R^2 statistic peaks at 14 quarters explaining 45% of the variation of foreign equity holdings in U.S. wealth. This is exactly what is suggested by the estimation of the error correction coefficients. However, theory would suggest that the highest predictive power should be displayed at business cycle frequency because of

[8] I focus on in-sample regressions because out-of sample regressions are not superior in terms of robustness in a setting like this (Inoue and Kilian (2004)). Furthermore, throughout the paper i calculate the cointegration residual with the cointegration coefficient estimates for the sample period from 1952 to 2004 and also use it for forecast regressions for shorter sample periods. Cointegration is a long-run relationship, that is why estimation of the cointegration coefficients only for a (shorter) forecast sample period would mean to throw away information.

time-varying risk premia over a business cycle[9], i.e. 10-12 quarters as is the case for total U.S. stock market wealth holdings (L&L 2004). Nevertheless, this result is in line with Fama and French (1988) and Poterba and Summers (1988) as well as Richards (1995) who showed that autocorrelation of expected stock returns is highest at four to five year horizons and hence returns should be best predictable at that frequency. Furthermore, it is a macroeconomically founded predictor that is regarded such that macroeconomic explanations for asset return predictability are still valid. Moreover, fluctuations of the market value of foreign equity holdings do not have to move with cyclical variation of the U.S. economy. The "home bias" of equity portfolios is a well established fact in the literature,[10] such that it seems to be reasonable that cyclical fluctuations of the U.S. economy will be first reflected in cyclical variation of risk premia for domestic stocks. U.S. households will react to the perception of cyclical fluctuations of the U.S. economy by requiring time-varying risk premia for U.S. stocks at business cycle frequency which leads to cyclical changes of the market value of their stock market wealth component as it is dominated by domestic stock holdings. Although there are tendencies of a syncronisation of business cycles for European countries[11], this does not imply that business cycles move together worldwide. In addition, there are idiosyncratic elements in national business cycles[12], which can be put forward to explain why changes of the market value of foreign equity holdings are not predictable at the same frequency as changes of U.S. households´ total stock market wealth are.

A temporary deviation of the market value of U.S. households´ foreign equity holdings is induced by the expectation of varying stock returns in the future. Time-varying risk premia over a business cycle could be responsible for that. Thus foreign stock returns vary over the foreign, national business cycles which do not necessarily move together with the U.S. business cycle. Hence, the best prediction horizon for changes of U.S. households´ foreign equity holdings does not have to coincide with U.S. business cycle frequency.

Another open question is how U.S. households allocate their foreign eq-

[9] Time-varying risk premia could be caused by habit formation (Campbell and Cochrane (1999)) or uninsurable background risks (Constantinidis and Duffee (1996), Heaton and Lucas (2000a,2000b))

[10] e.g. Tesar and Werner (1995)

[11] Artis and Zhang (1999)

[12] Artis et al. (1997)

uity portfolio geographically. This question encourages the assessment of the predictive power of $\widehat{\beta}' x_{t-1}$ for a rest-of-the-world stock market index which replaces the foreign equity holdings. As a first order approximation i assume that U.S. households predominantly invest in major stock markets for which information is widely available. Therefore i use a market capitalization weighted rest-of-the-world index constructed from data on the market capitalization of the MSCI stock indizes of the U.K., France, Italy, Canada, Japan and Germany, the G7 exclusive the U.S., in current U.S. dollars as an empirical proxy. The second column of Panel A presents forecast regressions of the changes of the rest-of-the-world index market capitalization with the cointegration residual as sole regressor for the sample period from fourth quarter 1969 to second quarter 2004. It can be obviously verified that the cointegration residual does not display any predictive power for this rest-of-the-world index at short (one to four quarter) horizon. However, it explains up to 28% of rest-of-the-world market capitalization changes at a time horizon of 14 quarters.

Panel B of table 2 offers a more detailed look on the predictive power of the cointegration residual. Here changes of the market capitalization of the individual MSCI stock indizes are regressed on the cointegration residual for the sample period from fourth quarter 1969 to second quarter 2004 in order to reveal what country indizes are particularly predictable and to what extent.

For the U.K. the cointegration residual displays predictive power at any forecast horizon and peaks at 14 quarters with an R^2 of 0.57. The forecasting ability for changes of the German MSCI stock market capitalization changes is also highest at the 14 quarter horizon but $\widehat{\beta}' x_{t-1}$ explains only 25% of the German stock market capitalization fluctuations. The highest predictive power for Canadian and French stock market capitalization changes is reached at 16 quarters, for the Italian MSCI stock index at 20 quarters. The Japanese stock market capitalization is not predictable at all as the regressor coefficients are not statistically different from zero at any time horizon. To summarize, with the exception of Japan, fluctuations of the stock market capitalizations of the MSCI stock indizes expressed in current U.S. dollars are highly predictable by the cointegration residual at 14 to 20 quarter frequency.

As stock market capitalizations are the underlying for index values, forecast regressions of returns on the MSCI indizes should confirm the results from panel B of table 2. The market capitalizations employed above are reported in current U.S. dollars which is of only interest for U.S. investors. In that context it would be interesting to know if the predictability of national stock market capitalizations displayed above is induced by movements of the stock market capitalization or changes of the respective bilateral exchange rate. From the U.S. point of view changes of $S_t P_t^{*,i}$ are predictable, with S_t, the bilateral U.S. dollar spot exchange rate at time t in British terms and $P_t^{*,i}$ the stock market capitalization of country i at time t in local currency. Predictability of $\Delta(S_t P_t^{*,i})$ could arise because ΔS_t or $\Delta P_t^{*,i}$ or both are predictable To assess this issue raw returns on the MSCI indizes in local currency as well as changes of the bilateral U.S. dollar spot exchange rate are regressed on the cointegration residual.

Table 3 reports results of the forecast regression of raw returns on the individual MSCI indizes in local currency regressed on the cointegration residual. The overall pattern is that $\widehat{\beta}' x_{t-1}$ is a strong predictor of movements of stock indizes in local currencies. It reaches its highest predictive power for 14 to 16 quarter returns well in accordance with the results for changes of the current dollar stock market capitalization changes with the exception of Italy. In detail, for the UK the peak of predictablility of the MSCI local currency index return is reached at 14 quarters explaining 51% of the return variation which is slightly lower than the R^2 for changes of the market capitalization in current dollars with an R^2 of nearly 57%. For France, $\widehat{\beta}' x_{t-1}$ predicts more than 54% of the local currency return on its MSCI index at 16 quarters horizon compared to 45% of the stock market capitalization changes in current dollars. Variation of the Italian real return on the respective MSCI index can be explained by the cointegration residual with an R^2 of 44% at 16 quarters. The R^2 is here slightly lower and peaks four quarters earlier than for the current dollar market capitalization changes. Canadian MSCI index returns are best predictable at 16 quarters with an R^2 of 31%. As for the changes of the MSCI market capitalization changes denominated in current dollars returns on the Japanese MSCI index in local currency are not predictable at all. All regressor coefficients are not significantly different from zero. Returns on the German MSCI index in domestic currency are best predictable at 14 quarters with an R^2 of 37% which is higher than the respective R^2 statistic of the regression presented in table 2.

Comparing the results of Panel B in table 2 and the regression outputs in table 3, it seems to be the case that irrespective of the exchange rate the cointegration residual explains movements of international stock markets.

This is mirrored by its explanatory power for returns on stock indizes with underlying market capitalization in local currency as well as for changes of market capitalizations in current U.S. dollars. The only exception is Japan for which no predictive power is revealed for both, market capitalization changes in current dollars and in yen. Furthermore, the changes of the R^2 statistic seem to display that fluctuations of the bilateral nominal exchange rates do not play an important role for forecasts of national stock market fluctuations. However, in the case of Canada, Italy and the U.K. the R^2 statistic of changes of the stock market capitalization in current dollars is higher than for returns on the local currency indizes. Hence, it could be possible that U.S. dollar exchange rate changes for these countries are predictable as well. Therefore changes of the bilateral nominal U.S. dollar exchange rate are regressed on the cointegration residual. As exchange rate the end of quarter bilateral spot rates published by the FRB are chosen. The sample for the EMU member countries France, Germany and Italy spans the first quarter 1971 to fourth quarter 1998 as reported by the FRB. Here changes of the French Franc, Italian Lira and Deutschmark are investigated. The time horizon for exchange rate changes of British Pound, Canadian Dollar and Japanese Yen covers the period from first quarter 1971 to second quarter 2004. Results are reported in table 4.

The overall picture that emerges is that nominal exchange rate changes are hardly predictable by $\widehat{\beta}' x_{t-1}$.Changes of the Lira are predictable at 16 and 20 quarter frequency explaining up to 15% of its variation. Movements of the U.S. Dollar - Lira exchange rate seem to be responsible for the change of the best prediction frequency of Italian stock market capitalization changes. At 16 and 20 quarters horizon $\widehat{\beta}' x_{t-1}$ explains 7% and 12%of the variation of the British Pound. But these are the only statistically significant regressor estimates in the forecast regressions on changes of the nominal exchange rate. Hence, the conclusion can be drawn that the nominal exchange rate plays only a minor role in forecasting international stock market movements reflected by changes of stock market capitalizations.

4 Final Remarks

Empirical evidence presented in this paper suggests that the U.S. consumption-wealth ratio embodies information on changes of foreign equity holdings of U.S. households and hence forecasts fluctuations of foreign stock market capitalizations. The residual of the cointegrating relation between non-durable and services consumption expenditure, foreign equity holdings of U.S. households, domestic asset wealth and labour income displays predictive power for changes of the market capitalization of foreign stock indizes at 14 to 20 quarter horizon. The influence of the nominal bilateral U.S. dollar exchange rate on the predictability of stock market capitalization changes is negligible. The main results shown in this paper underscore that international stock markets are linked which could help to further shed light on issues of international business cycle transmission through financial markets.

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A Data

- The definition of U.S. household stock market wealth includes directly held equity shares at market value and indirectly held equity shares namely bank personal trusts and estates holdings, life insurance companies´ holdings, private pension fund holdings, state and local government as well as federal government fund holdings and household´s mutual fund holdings as published in the supplemental table B.100e in the Z1 Flow of Funds Accounts of the Federal Reserve Board. However, this table is not available at quarterly frequency. That is why the value of quarterly stock market wealth is constructed with help of Flow of Funds tables L.213 and L.214 to match the values provided in table B.100e.

  - Table L.213 lists the direct holdings of corporate equity at market value distinguished by the respective holders. According to the definition above direct equity holdings of the household sector (line 6), bank personal trusts and estates (line 11), life insurance companies (line 12), private pension funds (line 14), state and local government (line 15) as well as federal government corporate equity holdings (line 16) are included. The amount of equities directly and indirectly held by U.S. households through mutual fund holdings is constructed with help of table L.214.
  - Table L.214 lists the direct holdings of mutual fund shares at market value distinguished by the respective holders. In order to calculate the amount of equities held by U.S. households through mutual fund holdings, the fraction of e.g. direct household mutual fund shares holdings at market value is calculated and multiplied with the direct holding of corporate equities by mutual funds

(L.213, line 17). This procedure is applied to all components of stock market wealth listed above which hold mutual fund shares and hence indirectly corporate equity.

- The share of foreign equity in household net worth is calculated with help of Flow of Funds table L.213 which provides details about equity issues and holdings at market value. Corporate equity issues at market value include holdings of foreign issues by U.S. residents inclusive American Depositary Receipts. It is assumed that the share of this rest-of-the-world equity holdings in total corporate equity holdings is the same as the share of rest-of-the-world equity holdings in U.S. households´ corporate equity holdings which is a reasonable approximation as U.S. households hold roughly 90% of total corporate equity issues.

- U.S. household domestic asset wealth is simply defined as difference between household net worth, Z1 flow of funds table B.100, line 42, and U.S. foreign equity holdings defined above.

- U.S. consumption is defined as consumption expenditure on non-durable goods and services excluding footwear and clothing published by the Bureau of Economic Analysis in NIPA table 2.3.5 and follows the definition used by L&L.

- Data on U.S. labour income is freely available from the Bureau of Economic Analysis in NIPA table 2.1. Labour income is defined as wages and salaries disbursements (line 3) + employer contribution for employee pension and insurance funds (line 7) + personal current transfer receipts (line 16) - contributions for government social insurance (line 24) - labour taxes. Labour taxes are defined as {wages and salaries disbursements / [wages and salaries disbursements + proprietors´ income with inventory valuation and capital consumption adjustment (line 9) + rental income of persons with capital consumption adjustment (line 12) + personal interest income (line 14) + personal dividend income (line 15)]} times [personal taxes (line 25) + personal current transfer payments (line 30)].

- Real variables are obtained by deflating with the CPI deflator of total personal consumption expenditure in chain-weighted (2000 = 100) seasonally adjusted U.S. dollars published by the Bureau of Economic Analysis in NIPA table 1.1.4.

- Per capita variables are obtained with population figures from NIPA table 2.1 published by the Bureau of Economic Analysis.

- The rest-of-the-world stock index is defined as value-weighted sum of the end of quarter market capitalization in current U.S. dollars of the Morgan Stanley Capital International stock indizes for the United Kingdom, France, Italy, Canada, Japan and Germany for which end of month data from December 1969 till June 2004 was provided by Morgan Stanley Capital International. Quarterly data is obtained by using end of period values.

- Changes of the current dollar MSCI stock index capitalizations are defined as natural logarithm of the market capitalization at time t+1 minus the natural logarithm of the market capitalization at time t. As logarithmic approximations of market capitalization changes are regarded the h-period market capitalization change is simply the sum of the one period market capitalization changes over h periods.

- Raw returns on MSCI indizes in local currencies for the G7 excluding the U.S. are defined as natural logarithm of the respective index value at time t+1 minus the natural logarithm of the index value at time t. As logarithmic approximations of returns are regarded the h-period return is simply the sum of the one period returns over h periods.

- Changes of the bilateral U.S. dollar spot exchange rates are obtained from daily spot exchange rate data published on the FRB webpage. The quarterly spot exchange rate is defined as end of quarter spot exchange rate. Changes of the exchange rate are defined as natural logarithm of the spot rate at time t+1 minus the natural logarithm of the spot rate at time t. As logarithmic approximations of spot rate changes are regarded the h-period spot rate change is simply the sum of the one period spot rate changes over h periods.

  For Germany, France and Italy the spot exchange rates of the U.S. dollar with Deutschmark, French Franc and Italian Lira are analysed for the time period from first quarter 1971 to fourth quarter 1998. For Canada, Japan and the U.K. spot exchange rates of the U.S. Dollar with Canadian Dollar, Japanese Yen and British Pound are investigated.

Table 1: Johansen Cointegration Test

r =	Trace test critical values 10 %	5%	1%	L-max test critical values 10 %	5%	1%
0	44.4929	47.8545	54.6815	25.1236	27.5858	32.7172
1	27.0669	29.7961	35.4628	18.8928	21.1314	25.8650
2	13.4294	15.4943	19.9349	12.2971	14.2639	18.5200
3	2.7055	3.8415	6.6349	2.7055	3.8415	6.6349

	Trace	L-max
1 lag	44.1243	25.0711
	19.0531	15.4426
	3.6106	3.5955
	0.0150	0.0150
2 lags	35.1189	19.0980
	16.0209	12.8401
	3.1808	3.1788
	0.0021	0.0021

	AIC	SIC
1 lag	-22.0350	-21.7791
2 lags	-21.9342	-21.4224

Notes: The variables employed are non-durables and services consumption excluding expenditures on footwear and clothing, foreign equity holdings of U.S. households, domestic asset wealth and labour income. All variables are measured at quarterly frequency. The sample period starts second quarter 1952 and ends second quarter 2004.
The Johansen test is performed under the assumption of an unristricted constant but no time trend in the data. The Trace test tests the null hypothesis of r cointegrating relations against the alternative of p, the number of variables in the tested system, cointegrating relations. The L-max test tests the null of r cointegrating relations against the alternative of r+1. AIC is the Akaike information criterion, SIC the Schwartz information criterion.

Table 2: Forecast Regressions

Panel A

h	ΔRoW	ΔRoW-Index
1	1.2885 ; R^2: 0.0702 (3.2733)	0.6258 ; R^2: 0.0175 (1.5367)
4	4.1966 ; R^2: 0.2278 (4.3917)	2.5906 ; R^2: 0.0791 (1.6913)
8	7.3138 ; R^2: 0.3426 (5.7321)	4.9717 ; R^2: 0.1299 (2.0329)
12	10.6119 ; R^2: 0.4439 (7.0916)	8.4411 ; R^2: 0.2526 (3.1510)
14	11.7825 ; R^2: **0.4495** (7.2716)	9.5055 ; R^2: **0.2826** (3.5419)
16	12.5687 ; R^2: 0.4272 (6.3476)	9.8514 ; R^2: 0.2693 (3.4848)
20	14.4614 ; R^2: 0.4006 (6.8969)	8.4574 ; R^2: 0.1669 (3.0201)

Panel B

h	ΔUK	ΔFRA	ΔITA
1	0.9353 ; R^2: 0.0363 (2.2442)	0.9134 ; R^2: 0.0222 (1.7628)	0.8854 ; R^2: 0.0150 (1.5319)
4	4.1324 ; R^2: 0.2026 (2.9685)	3.9512 ; R^2: 0.1037 (2.3249)	4.5163 ; R^2: 0.1018 (2.2235)
8	7.2059 ; R^2: 0.3496 (3.8481)	7.3177 ; R^2: 0.1629 (2.8210)	9.5061 ; R^2: 0.2000 (2.9338)
12	11.0587 ; R^2: 0.5279 (6.5690)	12.5026 ; R^2: 0.3409 (4.8816)	15.8403 ; R^2: 0.3851 (5.0323)
14	11.9687 ; R^2: **0.5678** (7.2627)	14.6858 ; R^2: 0.4145 (5.7572)	18.3000 ; R^2: 0.4551 (6.1434)
16	12.5252 ; R^2: 0.5454 (7.5270)	16.3825 ; R^2: **0.4511** (6.2210)	20.2442 ; R^2: 0.4927 (7.1195)
20	12.9532 ; R^2: 0.5180 (6.2874)	17.4426 ; R^2: 0.4417 (6.5490)	22.0814 ; R^2: **0.4971** (9.5473)

Table 2 (continued): Forecast Regressions

Panel B (continued)

h	ΔCND	ΔJPN	ΔGER
1	0.7403 ; R^2: 0.0217 (1.9509)	0.2146 ; R^2: -0.0057 (0.3479)	0.8379 ; R^2: 0.0205 (1.8268)
4	2.1870 ; R^2: 0.0526 (1.7821)	0.6750 ; R^2: -0.0046 (0.2779)	3.4057 ; R^2: 0.0976 (2.2088)
8	3.8312 ; R^2: 0.1182 (2.3637)	1.7278 ; R^2: 0.0004 (0.4322)	5.7727 ; R^2: 0.1362 (2.1213)
12	5.9953 ; R^2: 0.2815 (4.9809)	4.1232 ; R^2: 0.0238 (0.8766)	8.9824 ; R^2: 0.2459 (2.6228)
14	7.1841 ; R^2: 0.4030 (8.0883)	4.7951 ; R^2: 0.0310 (1.0130)	9.5228 ; R^2: **0.2543** (2.7462)
16	7.7464 ; R^2: **0.4365** (8.1991)	4.3318 ; R^2: 0.0207 (0.8915)	9.7649 ; R^2: 0.2396 (2.6806)
20	8.5200 ; R^2: 0.3868 (6.5726)	0.4987 ; R^2: -0.0080 (0.1080)	7.6462 ; R^2: 0.1238 (2.1516)

Notes: Table 2 reports OLS regression results with the cointegration residual as sole regressor. The forecast horizon h is in quarters. Panel A displays the results for forecasts of changes of the rest of foreign equity holdings in the U.S. household stock market wealth component, ΔRoW, and changes of the market capitalization of a value-weighted rest of the world index , ΔRoW-Index. The sample spans the period from second quarter 1952 to second quarter 2004 for forecasts of ΔRoW and the period from fourth quarter 1969 to second quarter 2004 for ΔRoW-Index. Panel B presents details on the predictive power of the cointegration residual on changes of the market capitalization of the individual MSCI stock indizes used to construct the value-weighted rest of the world index in Panel A. The sample covers quarterly observations from fourth quarter 1969 to second quarter 2004. R^2 reports values of the adjusted R^2. Newey-West corrected t-statistics for the significance of the regressor coefficient estimates are provided in parentheses.

Table 3: Forecast Regressions

h	rrUK	rrFRA	rrITA
1	1.0705 ; R^2: 0.0568 (2.6417)	0.8899 ; R^2: 0.0277 (2.2277)	0.7792 ; R^2: 0.0140 (1.6107)
4	4.1422 ; R^2: 0.2164 (3.0744)	3.5171 ; R^2: 0.1242 (2.6849)	3.7902 ; R^2: 0.0955 (2.3137)
8	7.1820 ; R^2: 0.3690 (4.2257)	6.4323 ; R^2: 0.2112 (3.5746)	7.9169 ; R^2: 0.1916 (3.2250)
12	9.9707 ; R^2: 0.4900 (5.7389)	10.7963 ; R^2: 0.4407 (6.0026)	12.8654 ; R^2: 0.3584 (5.8782)
14	10.5222 ; R^2: **0.5104** (6.0186)	12.5944 ; R^2: 0.5230 (7.3609)	14.8685 ; R^2: 0.4186 (7.4684)
16	10.4417 ; R^2: 0.4591 (5.6993)	13.7490 ; R^2: **0.5473** (8.2743)	16.4139 ; R^2: **0.4424** (8.6502)
20	9.5389 ; R^2: 0.3355 (4.7357)	14.5066 ; R^2: 0.5248 (8.2691)	17.8777 ; R^2: 0.4035 (8.9814)

h	rrCND	rrJPN	rrGER
1	0.4598 ; R^2: 0.0090 (1.4173)	0.2912 ; R^2: -0.0027 (0.6930)	0.9908 ; R^2: 0.0378 (3.0573)
4	1.6364 ; R^2: 0.0370 (1.4351)	0.7166 ; R^2: -0.0011 (0.4378)	3.5244 ; R^2: 0.1351 (3.3914)
8	3.1159; R^2: 0.0907 (2.0542)	1.5372 ; R^2: 0.0061 (0.6180)	5.7930 ; R^2: 0.1891 (3.1049)
12	5.2941 ; R^2: 0.2257 (4.3206)	3.0349 ; R^2: 0.0279 (1.0357)	9.0287 ; R^2: 0.3395 (3.5518)
14	6.1315 ; R^2: 0.2926 (5.8108)	3.4377 ; R^2: 0.0320 (1.1508)	9.8288 ; R^2: **0.3706** (3.9290)
16	6.4356 ; R^2: **0.3098** (6.2460)	3.3876 ; R^2: 0.0251 (1.0678)	10.2346 ; R^2: 0.3659 (4.0515)
20	6.6668 ; R^2: 0.2652 (5.0548)	1.4936 ; R^2: -0.0039 (0.4706)	9.8250 ; R^2: 0.2915 (4.0242)

Notes: Table 3 reports OLS regression results of the real return on MSCI stock indizes with underlying market capitalization in domestic currency. The cointegration residual is the only regressor. The forecast horizon h is in quarters. The sample covers quarterly observations from fourth quarter 1969 to second quarter 2004. R^2 reports values of the adjusted R^2. Newey-West corrected t-statistics for the significance of the regressor coefficient estimates are provided in parentheses

Table 4: Forecast Regressions

h	ΔS_{UK}	ΔS_{FRA}	ΔS_{ITA}
1	0.0277 ; R^2: -0.0075 (0.2645)	-0.2160 ; R^2: -0.0018 (-0.9384)	-0.1111 ; R^2: 0.0140 (-0.5956)
4	0.4616 ; R^2: 0.0029 (1.2224)	-0.3854 ; R^2: -0.0054 (-0.4724)	0.2346 ; R^2: -0.0076 (0.3390)
8	0.7871 ; R^2: 0.0061 (1.0650)	-0.5460 ; R^2: -0.0057 (-0.4558)	0.9275 ; R^2: 0.0026 (0.7712)
12	1.7593 ; R^2: 0.0405 (1.5846)	0.2639 ; R^2: -0.0096 (0.1842)	2.6540 ; R^2: 0.0564 (1.7903)
16	2.4938 ; R^2: 0.0734 (2.1112)	1.0539 ; R^2: -0.0028 (0.7283)	3.8705 ; R^2: 0.1019 (2.6988)
20	3.2311 ; R^2: **0.1232** (2.9405)	1.5958 ; R^2: 0.0038 (0.9373)	5.1594 ; R^2: **0.1497** (8.0869)
h	ΔS_{CND}	ΔS_{JPN}	ΔS_{GER}
1	0.0061 ; R^2: -0.0076 (0.0633)	0.0146 ; R^2: -0.0076 (0.6930)	-0.4407 ; R^2: 0.0181 (-1.9546)
4	-0.1116 ; R^2: -0.0050 (-0.4672)	0.4528 ; R^2: -0.0004 (0.5629)	-0.9354 ; R^2: 0.0190 (-1.2729)
8	-0.4495; R^2: 0.0135 (-1.4174)	1.1250 ; R^2: 0.0135 (0.8504)	-1.3925 ; R^2: 0.0184 (-1.0343)
12	-0.8002 ; R^2: 0.0416 (-1.7618)	2.1687 ; R^2: 0.0506 (1.4382)	-1.0905 ; R^2: 0.0026 (-0.7226)
16	-0.6301 ; R^2: 0.0160 (-1.0486)	2.0301 ; R^2: 0.0356 (1.2850)	-0.7582 ; R^2: -0.0053 (-0.6515)
20	-0.0627 ; R^2: -0.0087 (-0.0772)	0.2312 ; R^2: -0.0084 (0.1466)	-1.1083 ; R^2: -0.0017 (-0.6660)

Notes: Table 4 reports OLS regression results of bilateral US-dollar spot exchange rate changes with the cointegration residual as sole regressor. The forecast horizon h is in quarters. The sample for EMU member countries covers quarterly observations from first quarter 1971 to fourth quarter 1998. The sample for the three other countries covers the period from first quarter 1971 to second quarter 2004. R^2 reports values of the adjusted R^2. Newey-West corrected t-statistics for the significance of the regressor coefficient estimates are provided in parentheses