The U.S. consumption-wealth ratio and foreign stock markets: International evidence for return predictability

Thomas Nitschka[1]
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 builds on the approach developed by Lettau and Ludvigson (2001 JoF, 2004 AER) to investigate if the U.S. consumption-wealth ratio is not only informative about the future path of U.S. but also foreign stock markets.

Evidence presented in this paper suggests that a four-variable logarithmic approximation of the U.S. consumption-wealth ratio cointegrates and embodies information about changes of the market value of U.S. households´ foreign stock holdings. Variations of U.S. households´ foreign equity holdings are predominatly driven by the transitory shock in the cointegrated system under consideration. A VECM analysis displays that a deviation from the cointegration trend among consumption and wealth is adjusted through temporary fluctuations of foreign equity at market value. Consistent with that finding it is shown that the respective cointegration residual is a good predictor of excess returns on foreign stock indizes at business cycle frequency. Exchange rate changes seem to play a negligible role in this context.

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 framework proposed by Lettau and Ludvigson (2001, 2004) 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) provide evidence that the consumption-wealth ratio cointegrates in U.S. data. They find that the cointegration residual of their trivariate empirical approximation of the log consumption-wealth ratio 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. The highest predictive power is reached at business cycle frequency which seems to be driven by time-varying risk premia over a business cycle. Fernandez- Corugedo et al. (2003), Fisher and Voss (2004) and Tan and Voss (2003) corroborate these findings in U.K. and Australian data. Hamburg et al. (2005) find one cointegration relation among a trivariate German consumption-wealth ratio approximation. However, deviations from the common trends are adjusted by temporary fluctuations of disposable income which seems to be the outcome of structural differences between the German and Anglo-Saxon financial system. Hence, the German consumption-wealth ratio does not display predicitve power for German stock market returns but business cycle variables such as the unemployment rate.

Apart from the special case of Germany, consumpion-wealth ratios of Anglo-Saxon countries mirror movements of national stock markets. What, to my knowledge, has not been done so far is to assess the question if a national consumption-wealth ratio does not only mirror fluctuations of national but also foreign stock markets. The motivation for this question is straightforward. L&L (2004) show that transitory movements of U.S. stock market wealth, i.e. equity holdings of U.S. households, are primarily responsible for temporary fluctuations of the consumption-wealth ratio and hence cause the predictive power of the cointegration residual for returns on broad U.S.

[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.

stock indizes. As U.S. households either directly or indirectly hold foreign equity, the U.S. consumption-wealth ratio should also embody information about movements of foreign stock markets U.S. households are invested in. I focus on the U.S. consumption-wealth ratio given the importance of the U.S. economy and the importance of the absolute value of U.S. households´ foreign equity investment. Furthermore, I suppose that foreign equity holdings at market value are primarily driven by the transitory shock in the cointegration relation between consumption and wealth in U.S. data such that temporary fluctuations of the consumption-wealth ratio reflect expected returns on (changes of) U.S. households´ foreign equity holdings at market value.

Evidence presented in this paper suggests that a four-variable logarithmic approximation of the U.S. consumption-wealth ratio cointegrates. The consumption-wealth ratio proxy consists of non-durable and services consumption expenditure excluding clothing and footwear, U.S. households´ foreign equity holdings, domestic asset wealth and labour income. The error correction representation of this cointegrated system with one cointegration relation allows to infer that foreign equity holdings adjust a deviation from the common trends. Moreover, the impact of permanent and transitory shocks in this cointegrated system on consumption, foreign equity, domestic asset wealth and labour income is quantified and clearly reveals that foreign equity holdings are predominantly driven by the transitory shock in the cointegrated system under consideration. It is shown that the cointegration residual is not only a good predictor of changes of foreign equity holdings at market value but also of excess returns on foreign stock indizes at business cycle frequency irrespective if the underlying market capitalization is denominated in current U.S. dollars or local currency. Hence, changes of U.S. dollar exchange rates seem to play a negligible role in this context.

The remainder of this paper is organized as follows. A simple manipulation of the theoretical framework used by L&L as well as the cointegration properties of the four-variable consumption-wealth ratio proxy are discussed in section two. Section three provides details on forecasts of changes of U.S. households´ foreign equity holdings and excess returns on foreign stock indizes. 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 trends 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 small relative to total wealth but I suppose given the importance of the U.S. economy and the importance of the absolute value of these holdings that 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 model 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 wealth 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 natural logarithms of (1) gives (2). In the following lower-case letters denote logarithms.

$$\begin{aligned} w_{t+1} - w_t &= r_{w,t+1} + \log(1 - \frac{C_t}{W_t}) && (2) \\ w_{t+1} - w_t &= r_{w,t+1} + \log(1 - \exp(c_t - w_t)) && (3) \end{aligned}$$

Under the assumption that the log consumption-wealth ratio is covariance stationary and a long-run mean exists, then the last term of the right-hand side of (3) can be approximated via a Taylor expansion around the mean of the consumption-wealth ratio, $\overline{c-w}$. 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 $w_{t+1} - w_t \equiv \Delta w_{t+1}$ one obtains

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

Writing Δw_{t+1} tautologically in terms of the consumption growth rate and changes of the consumption-wealth ratio

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

as well as substitution in (5) yields

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

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 law of motion of 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 wealth, is unobservable.

This issue can be solved by assuming that aggregate labour income is the dividend paid from human wealth and represents its non-stationary component. Define the gross return on human capital 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 wealth 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 wealth is the present discounted value of expected labour income. Returning to the one-period scenario and employing the Campbell-Shiller return decomposition (Campbell and Shiller (1988)) under the assumption that labour income as well as human wealth are integrated of order one, I(1), and additionally assuming that the ratio of log labour income and log human wealth is covariance stationary 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 wealth 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

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

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

if expectations are taken on both sides. Exploiting the assumption that labour income is I(1) equation (13) gives an expression of human wealth 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, decompose aggregate wealth into its components asset and human wealth

$$W_t = A_t + H_t$$

with A_t representing asset wealth and H_t human wealth. Assume that the ratio of human to asset wealth is covariance stationary and hence a long-run mean of this ratio exists. Then log aggregate wealth can be approximated around this mean which leads to

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

with v interpretable as average share of asset wealth in aggregate wealth. Furthermore, employing the same technique, asset wealth at time t, A_t, can be decomposed into its components foreign equity, FE_t, and the rest of asset wealth which I will refer to as domestic asset wealth, DAW_t, such that

$$A_t = FE_t + DAW_t$$

Thus the logarithmic approximation of asset wealth around the long-run mean of the covariance stationary foreign equity to domestic asset wealth ratio obeys

$$a_t \approx \lambda fe_t + (1-\lambda)daw_t \tag{15}$$

with λ the average share of foreign equity in U.S. households´ asset wealth.

A combination of (14) and (15) then gives

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

with $\theta = v\lambda$ the average share of foreign equity in total wealth, $\phi = v(1-\lambda)$ the average share of domestic asset wealth in aggregate wealth.

Additionally, the gross return on aggregate wealth is decomposed into the returns on its components, first into the components asset and human wealth which yields

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

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

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

Plugging (19) and (16) into (7) and taking expectations on both sides of the equation gives

$$\begin{aligned} & c_t - \theta fe_t - \phi daw_t - (1 - \theta - \phi) h_t \\ = \; & E_t\{\sum_{i=1}^{\infty} \rho_w^i [(\theta r_{fe,t+i} + \phi r_{daw,t+i} + (1 - \theta - \phi) r_{h,t+i}) - \Delta c_{t+i}]\} \end{aligned} \tag{20}$$

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 $\rho_w^i = \rho_h^i$.

$$\begin{aligned} & c_t - \theta fe_t - \phi daw_t - (1 - \varphi - \phi) y_t \\ = \; & E_t\{\sum_{i=1}^{\infty} \rho_w^i [(\theta r_{fe,t+i} + \phi r_{daw,t+i} + (1 - \theta - \phi) \Delta y_{t+i}) - \Delta c_{t+i}] + (1 - \theta - \phi) z_{t+i}\} \end{aligned} \tag{21}$$

According to (21), c_t, log consumption, fe_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 temporary deviation from the common trends, should either mirror changes of (returns on) foreign equity holdings, changes of domestic asset wealth, changes of labour income or consumption growth, or an arbitrary combination.

2.2 Empirical evidence: Cointegration and error correction

This section assesses the question if the four-variable proxy of the log consumption-wealth ratio proposed above cointegrates. All variables employed are quarterly, real, per capita, expressed in billions of chain-weighted 2000 U.S. 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. 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. 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.[2] 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. 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.[3] Moreover, estimates of

[2] Results available upon request

[3] 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.

the cointegration vector are highly plausible with respect to theory, which is addressed below. That is why I am convinced that cointegration among non-durables and services consumption, foreign equity holdings, domestic asset wealth and labour income exists and base the subsequent analyses on this assumption.

One reason to assume 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 first differences of foreign equity holdings, domestic asset wealth and labour income. Hence the estimate equation takes the following form

$$\begin{aligned} c_t &= \alpha + \beta_{fe} fe_t + \beta_{daw} daw_t + \beta_y y_t \\ &\quad + \sum_{i=-k}^{k} b_{fe,i} \Delta fe_{t-i} + \sum_{i=-k}^{k} b_{daw,i} \Delta daw_{t-i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t-i} + \varepsilon_t \end{aligned} \tag{22}$$

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.[4]

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

At first glance the estimated cointegration coefficients of foreign equity holdings, domestic asset wealth and labour income do not seem to be economically meaningful as they sum to a number bigger than unity. However,

[4] The estimates do not vary much from one to seven leads and lags. Here six leads and lags are employed. Johansen´s maximum likelihood procedure provides very similar estimates.

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 larger than one. Estimation of the cointegration vector using total personal consumption instead of only non-durables and services consumption expenditure leads to the following estimates:

$$\widehat{\beta} = [1 \; - \underset{(2.8860)}{0.0132} fe_t \; - \underset{(8.9216)}{0.2296} daw_t \; - \underset{(25.2976)}{0.7684 y_t}]'$$

The sum of these cointegration coefficients is approximately unity, to be precise it is 1.0112. However, for reasons mentioned before I use non-durables and services consumption expenditure as preferred proxy of consumption. In that case the sum of the cointegration coefficients increases by 8% to 1.0846. An interpretation of this finding is that the present value of durable consumption amounts to 8% of the present value of total consumption. To assess this point I investigated what the share of durable goods in household net worth is. The Federal Reserve Board of Governors reports replacement costs of durable goods in household net worth data. According to that data set the share of durable goods in household net worth is around 8% on average over time. Hence, the estimated cointegration vector using only non-durable consumption expenditures should reveal that in the long-run the net present value of asset wealth including the stock of durable goods should be about 8% higher than the present value of non-durable consumption. This is roughly consistent with the respective cointegration coefficient estimates when non-durable and services consumption expenditure are employed.

Furthermore, assuming that aggregate wealth represents output governed by a Cobb-Douglas production function, then the cointegration coefficients could be interpreted as reflecting the average shares of capital and labour in output which are stable over time. The share of labour would be approximately 0.7, the share of capital 0.3. A number close to values employed in the real business cycle literature[5] which additionally corroborates the findings by L&L (2001,2004). The point estimate of the cointegration coefficient of foreign equity holdings seems to be reasonable as well, i.e. it mirrors the average share of foreign equity in total wealth. The share of foreign equity in

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

U.S. households´ asset wealth considerably increased since the late 1980s 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 $\mathbf{x}_t = (c_t, fe_t, daw_t, y_t)'$ is

$$\mathbf{\Gamma}(\mathbf{L})\mathbf{\Delta x}_t = \boldsymbol{\alpha}\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{24}$$

where $\mathbf{\Delta x}_t = (\Delta c_t, \Delta fw_t, \Delta daw_t, \Delta y_t)'$ is the vector of first differences and $\mathbf{x}_{t-1}$ the vector of lagged levels, $\boldsymbol{\alpha} = (\alpha_c, \alpha_{fw}, \alpha_{dw}, \alpha_y)'$ is the vector of adjustment coefficients which reflect what variables are responsible for the error correction. $\mathbf{\Gamma}(\mathbf{L})$ denotes a (4x4) matrix polynomial in the lag operator and $\widehat{\boldsymbol{\beta}} = (1, -\widehat{\beta}_{fe}, -\widehat{\beta}_{daw}, -\widehat{\beta}_y)'$ represents the vector of the above estimated cointegration coefficients. Hats indicate estimated variables and $\boldsymbol{\varepsilon}_t$ represents the (4x1) vector of shocks in the cointegration relation with covariance matrix $\mathbf{\Omega}$. Lower-case letters in bold face denote vectors, bold upper-case letters represent matrices.

The term $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ gives the cointegration residual, $\boldsymbol{\alpha}$ is the adjustment vector that displays what variables adjust a deviation from the common trend among consumption and wealth. This is one conclusion that can be drawn from the Granger Representation Theorem: If $\mathbf{x}_t$ is cointegrated, at least one of the adjustment coefficients $\alpha_c, \alpha_{fe}, \alpha_{dw}$ or α_y must be nonzero in the error-correction representation.

All VECM coefficients are estimated by OLS applying a lag length of one 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.[6] T-statistics of the adjustment coefficient

[6] Results of the estimation of the remaining VECM coefficients are available from the author upon request. However, the main findings in L&L (2001) are corroborated.

estimates are reported in parentheses.

$$\boldsymbol{\alpha} = (\underset{(-0.8541)}{-0.0118}, \underset{(4.3329)}{1.2986}, \underset{(3.7045)}{0.2252}, \underset{(-0.1429)}{-0.0043})'$$

Apparently both asset wealth components are responsible for a restoration to the common trend. Domestic asset wealth adjusts to the common trend among consumption and total wealth which is persumably driven by the domestic stock market wealth component according to L&L (2004). Furthermore, the foreign equity holdings adjustment coefficient is not only relatively high but also statistically significant. Hence the conclusion can be drawn that a temporary variation of the U.S. consumption-wealth ratio contains information about 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. An interesting question that arises in this context is what the adjustment coefficient of foreign equity holdings exactly mirrors. It is not straightforward to see how the adjustment of a temporary deviation from the common trend among consumption and wealth through foreign equity holdings is achieved. One interpretation is that U.S. households rebalance their equity portfolio with respect to expected returns on foreign equity. However, as the market value of foreign equity holdings used in the VECM estimation is expressed in U.S. dollar terms, it could also be the case that the adjustment coefficient mirrors relatively stable quantities of foreign equity holdings over time, but fluctuations of the U.S. dollar exchange rate vis-á-vis the rest of the world or variations of foreign equity prices in local currency or an arbitrary combination.

2.3 Identification of permanent and transitory shocks and variance decomposition

As further device for the robustness of the previous results I follow Hoffmann (2001) to identify permanent and transitory shocks in the cointegrating system to quantify their contribution to the forecast error variance of the levels of the four cointegrated variables consumption, foreign equity, domestic asset wealth and labour income.

Since I regard a cointegrated system with four variables and one single cointegration relation there are three permanent shocks representing the innovations to the three common trends and one single transitory shock (Stock and Watson (1988)). Identification is achieved by inverting the vector error correction representation of $\mathbf{x}_t = (c_t, fe_t, daw_t, y_t)'$ into a multivariate Beveridge-Nelson moving average representation in terms of the reduced form disturbances (Beveridge and Nelson (1981)) which is given by

$$\mathbf{x}_t = \mathbf{C}(\mathbf{1}) \sum_{i=0}^{t} \varepsilon_i + \mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t \tag{25}$$

$\mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t$ denotes the stationary part of the moving average representation of $\mathbf{x}_t$ and $\mathbf{C}(\mathbf{1}) \sum_{i=0}^{t} \varepsilon_i$ represents the random-walk component.

Johansen (1995) has shown that $\mathbf{C}(\mathbf{1})$ can be identified with the parameters of the VECM such that

$$\mathbf{C}(\mathbf{1}) = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma}(\mathbf{1})\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp} \tag{26}$$

where $\boldsymbol{\beta}_{\perp}, \boldsymbol{\alpha}_{\perp}$ are the orthogonal complements of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. The Granger representation theorem implies that $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ satisfy $\boldsymbol{\beta}'\mathbf{C}(\mathbf{1}) = 0$ and $\mathbf{C}(\mathbf{1})\boldsymbol{\alpha} = 0$. Thus the common trends are

$$\pi_t = \boldsymbol{\alpha}'_{\perp} \sum_{i=0}^{t} \varepsilon_i = \sum \eta_t. \tag{27}$$

Let $\boldsymbol{\eta}_t^P = \boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ denote the permanent shocks to the cointegrating relation and $\eta_t^T = \boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1}$ the transitory shock if it is orthogonal to the permanent shocks. Hence the structural permanent shocks and the structural transitory shock can be identified via the relation

$$\boldsymbol{\eta}_t = \mathbf{S}\boldsymbol{\varepsilon}_t \tag{28}$$

with $\boldsymbol{\eta}_t = \binom{\eta_t^P}{\eta_t^T}$ and $\mathbf{S} = \binom{\boldsymbol{\alpha}'_{\perp}}{\boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1}}$ requiring that η_t^P and η_t^T have unit variance.

With this identification it is straightforward to quantify the contribution of the combined permanent shocks and the single transitory shock to the forecast error variance of the four cointegrated variables. Table 2 presents the decomposition of the forecast error variance of the levels of c, fe, daw and y into the components that can be attributed to the three permanent shocks combined and to the transitory shock. It is assumed that the transitory shock is orthogonal to the permanent shocks. The top panel reports the variance decomposition if statistically insignificant adjustment coefficient estimates are set to zero. The bottom panel displays the variance decomposition if all adjustment coefficients are set to their estimated values.

From the estimation of the adjustment coefficients it is clear that the transitory shock should have the strongest effect on the forecast error variance of both asset wealth components because their adjustment coefficient estimates are statistically significant. This implies that both participate in the correction of a temporary deviation from the common trends among c, fe, daw and y and hence should be primarily driven by the transitory shock. I.e., the transitory shock is mainly responsible for variations of fe and daw. The variance decompositions mirror exactly this reasoning. Note also that the impact of the transitory shock on the variance of foreign equity is stronger than on domestic asset wealth which is totally in line with the magnitude of the coefficient estimates. The foreign equity adjustment coefficient is substantially larger than that of domestic asset wealth, i.e. the transitory shock has to have a stronger impact on foreign equity than on domestic asset wealth. The opposite reasoning applies for consumption and labour income whose adjustment coefficients are statistically indistinguishable from zero, which means that both variables participate little in the error-correction mechanism and hence should be predominantly driven by the permanent shocks. The variance decompositions for consumption and labour income support this reasoning as well. Almost all of the variation of consumption and labour income can be attributed to the permanent shocks at any time horizon. This finding further corroborates the results of L&L (2004).

2.4 High turnover of foreign equity and the U.S. household equity portfolio

Though national equity portfolios are strongly biased towards home equity, as is the equity portfolio of U.S. households, data on cross-border equity flows reveals that the turnover rate of foreign equity is higher than the turnover rate of domestic equity. It seems to be the case that frequent transactions of institutional investors as mutual funds or pension funds are responsible for that finding rather than transactions generated by small, individual investors (Tesar and Werner (1995), Dahlquist and Robertsson (2001)). As on average over the sample period from second quarter 1952 to second quarter 2004 U.S.

households indirectly hold about 30% of their equity through e.g. mutual fund shares or life insurance companies, it would be interesting to see if the observation of Tesar and Werner of a high turnover in foreign equity compared to the turnover rate in domestic, i.e. U.S., equity is reflected in the U.S. household equity portfolio.

I conjecture that a higher turnover rate of U.S. households´ foreign equity holdings compared to their U.S. equity holdings would mean that foreign equity adjusts a deviation from the common trend among consumption and wealth faster than U.S. equity does. A temporary variation of the consumption-wealth ratio is induced by the expectation of returns on U.S. households´ stock market wealth (L&L (2004)), i.e. returns on their equity portfolio, which presumably causes a rebalancing of U.S. households´ equity portfolio and hence creates turnover. Thus a higher turnover rate of foreign relative to U.S. equity should be associated with a quicker adjustment of a temporary deviation from the common trend among consumption and wealth through foreign equity than through U.S. equity holdings. Quicker adjustment is tantamount to saying that the transitory shock in the cointegration relation should have a stronger impact on foreign than on U.S. equity, i.e. the transitory shock in the cointegration relation between consumption and wealth should contribute more to the forecast error variance of foreign equity than to the forecast error variance of U.S. equity. A variance decomposition of U.S. equity and foreign equity is a way to assess this point. Table 3 provides the results of this exercise for the sample period from second quarter 1952 to second quarter 2004. The contribution of the transitory shock to the forecast error variance of U.S. equity is higher than to the forecast error variance of foreign equity at any time horizon. Hence, the observation of a relative high turnover rate of foreign equity does not seem to be mirrored in the U.S. household equity portfolio over the regarded time horizon.

3 Forecasting power of the cointegration residual

As can be easily inferred from the adjustment coefficients of the VECM and the variance decompositions in the previous section, the cointegration residual should serve as a predictor of changes of U.S. households´ foreign equity holdings. Before describing the evidence of various long-horizon regressions reported in tables 4 to 6 it may be useful to provide some economic intuition of what should be reflected in the regression outcomes. According to (7), a temporarily high consumption-wealth ratio should either mirror high expected returns on aggregate wealth or low expected consumption growth. Regarding the evidence presented in the previous section foreign equity and domestic asset wealth are responsible for transitory deviations from the common trend among consumption and wealth and hence should be predictable by the cointegration residual. L&L (2001,2004) have shown that consumption and labour income growth are not predictable but changes of asset wealth are. A high consumption-wealth ratio, i.e. a positive deviation from the cointegration trend should thus be associated with the expectation of higher future returns on U.S. households´ asset wealth, here decomposed into returns on foreign equity and domestic asset wealth. This should be reflected in positive regressor estimates in the forecast regressions.[7]

A theoretical explanation for a temporary variation of expected returns on foreign equity and domestic asset wealth could be time-varying risk premia induced by variations of risk aversion over a business cycle. Reasons for time-varying risk aversion of agents at business cycle frequency could be the formation of consumption habits (Campbell and Cochrane (1999)) or the presence of uninsurable background risks (Constantinides and Duffee (1996), Heaton and Lucas (2000a,2000b)). The theoretical framework of this paper allows for both and further explanations.

However, the main purpose of this paper is to assess the question if the U.S. consumption-wealth ratio is informative about the future path of foreign stock markets. So, U.S. households´ foreign equity investment is the variable of particular interest in this empirical exercise. Regarding the left column of table 4 it is easily verified that $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ is a powerful predictor of changes of U.S. households´ foreign equity holdings, Δfe. 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 and the variance decomposition.

[7] Throughout the paper 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)). I.e., I calculate the cointegration residual with the cointegration coefficient estimates for the whole sample period from 1952 to 2004 and use it in 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.

However, foreign equity holdings are denominated in current U.S. dollars. Predictability then means that changes of the quantity of foreign equity, changes of the price of foreign equity in the respective local currency or changes of the nominal U.S exchange rate or an arbitrary combination are responsible for a transitory deviation from the common cointegration trend among consumption and aggregate wealth and hence predictable. In order to answer this question I constructed a foreign equity investment weighted effective exchange rate of the U.S. dollar relative to the countries the U.S. hold equity of. I focused on countries in which the U.S. invest at least one percent of their foreign equity investment.[8] All in all the seventeen countries used in this analysis represent about 80% of U.S. foreign equity investments. Data is taken from the IMF´s coordinated portfolio survey 2001. I used the share of U.S. equity investment into a particular country from total U.S. foreign equity investment as a weight to construct the above mentioned effective exchange rate and assume that these weights, derived from 2001 data, are constant over the whole sample period. This assumption certainly biases the effective exchange rate towards the foreign equity investment pattern of the U.S. in recent years. However, I think this can be justified by considering that foreign equity holdings constituted a negligible part of U.S. asset wealth at the beginning of the sample period until the 1980s. Since then the share of foreign equity in asset wealth increased considerably and hence gained more importance in U.S. households´ asset portfolio. Taking this reasoning into account and achknowledging that I am focused on the effective U.S. dollar exchange rate in foreign equity investment, the weights described above should be sufficient to approximate the true foreign equity investment weighted U.S. dollar exchange rate.

All weights are rescaled so that they sum to unity. If exchange rate changes of the US-dollar relative to the countries the U.S. hold equity of are responsible for temporary fluctuations of foreign equity holdings, then changes of the equity investment weighted exchange rate should be forecastable with the cointegration residual. Furthermore, I investigated if changes of foreign equity holdings denominated in a weighted basket of national currencies are predictable. I employed the equity investment weighted exchange rate to obtain foreign equity holdings in such a compound currency. Pre-

[8] I omitted equity investment in offshore equity markets as Bermuda or Cayman Islands and concentrated on Australia, Canada, Finland, France, Germany, Hong Kong, Ireland, Italy, Japan, Korea, Mexico, Netherlands, Sweden, Singapore, Spain, Switzerland and the United Kingdom.

dictability of changes of these holdings would either reflect variation of equity prices in local currency and/or changes of the quantity of foreign equity holdings or both. However, the importance of these effects are indistinguishable in this exercise.

The middle column of table 4 displays evidence that changes of the effective exchange rate, $\Delta neer$ are not predictable by the cointegration residual for the time period from first quarter 1957 to third quarter 2003. None of the regressor coefficients are statistically distinguishable from zero. According to that finding, changes of foreign equity holdings in local currency, Δfe^{NC} , should be predictable by $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ which can be confirmed regarding the regression results in the right column of table 4. The peak of predictability is reached after 12 quarters explaining 43% of foreign equity holdings variation, slightly lower than the highest predictive power in the case of foreign equity in U.S. dollars but at the same order of magnitude.

The forecast regressions reported in table 4 reveal that the cointegration residual displays information about variations of foreign equity holdings which should reflect movements of foreign stock markets presumably induced by time-varying risk premia. The non-predictability of exchange rate changes conveys the notion that the correction of the cointegration error through foreign equity holdings is mainly induced by the expectation of returns on foreign equity and subsequent portfolio rebalancing. Exchange rate returns seem to play a minor role in the error correction mechanism. Thus, $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ should predict (excess) returns on broad foreign stock indizes which should reflect cyclical variation of risk premia irrespective if the underlying market capitalization is denominated in U.S. dollar or local currency. In order to assess this point I investigate if $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ reveals predictive power for excess returns on Morgan Stanley Capital International (MSCI) stock indizes for the countries employed to calculate the effective exchange rate. Table 5 reports forecast regression results for excess returns on these indizes with underlying market capitalization in U.S. dollars. The sample covers the fourth quarter 1969 to second quarter 2004, except for Finland, Ireland, Korea and Mexico. The sample period for Finland spans the period from first quarter 1982 to second quarter 2004, the sample period for Ireland, Korea and Mexico covers the first quarter 1988 to second quarter 2004. Definitions of excess returns are provided in the appendix.

The regression results of table 5 mirror that the cointegration residual predicts excess returns with underlying market capitalization in U.S. dollars best at 8 to 24 quarter frequency. With the exception of Japan, Mexico and Singapore which are not predictable at any time horizon all MSCI stock index returns in U.S. dollars are to some extent predictable. A notable outlier is Korea, which displays some forecastability at long horizons, 12 to 24 quarters, but the regressor coefficient is negative, quite in contrast to the economic intution given at the beginning of the section. However, an explanation for this is straightforward. Korea experienced a severe currency crisis which had significant negative impact on the Korean stock market, while the U.S. enjoyed the stock market boom of the late 1990s right at the time of the Korean currency crisis which leads to the negative regressor coefficient estimates. A high U.S. consumption-wealth ratio is thus associated with negative returns on the Korean stock market.

Table 6 displays estimates from forecast regressions of excess returns on MSCI stock indizes with underlying market capitalization in local currency. The overall picture that emerges is that excess returns in national currency are in most cases even better predictable than returns in U.S dollars as can be seen from the magnitude of the R^2 statistics. It seems to be the case that exchange rate changes introduce noise and lower the explanatory power of the cointegration residual for stock index returns. The different peaks of predictability convey the notion that excess returns on stock indizes vary with the national business cycles which do not necessarily have to coincide with the U.S. cycle in frequency or magnitude because of idiosyncratic elements in national business cycles[9], although there are tendencies towards a syncronisation of business cycles for European countries[10]. But this does not imply that business cycles move together worldwide. This result is also in line with Richards (1995) who showed that the time horizon for which the highest autocorrelation of stock returns is reported differs internationally.

However, the frequency of the peak of predictability from the forecast regressions for U.S. households´ foreign equity holdings, i.e. 14 quarters when expressed in U.S. dollars or 12 quarters when expressed in compounded national currencies is almost identical with the peak of predictability for changes of total U.S. household stock market wealth which is reached at 12 quarters (L&L (2004))[11]. Thus it seems to be the case that U.S. households

[9] Artis et al. (1997)

[10] Artis and Zhang (1999)

[11] This is in line with Fama and French (1988a,1988b) and Poterba and Summers (1988) who provide evidence that the autocorrelation of stock returns is highest at three to five year horizon and thus returns should be best predictable at that frequency.

react to the perception of cyclical fluctuations of the U.S. economy with the adjustment of their equity portfolio at U.S. business cycle frequency irrespective if U.S. or foreign equity is regarded. Business cycle movements induce cyclical variations of U.S. households´ risk averison which leads to time-varying risk premia for U.S. stocks at U.S. business cycle frequency. Time-varying risk premia mirrored in excess returns cause a rebalancing of their equity portfolio. It could be the case that required risk premia for U.S. stocks are projected on foreign stocks because of the dominance of U.S. equity holdings in the U.S. households´ equity portfolio[12], such that their total equity portfolio is rebalanced at U.S. business cycle frequency. An alternative explanation for the coincidence of the peak of predictability of U.S. households´ foreign and total stock market wealth could be simply that U.S. households weight their foreign equity portfolio in such a way that it is positively correlated with the U.S. business cycle. An answer why U.S. households may do so is beyond the scope of this paper.

4 Conclusions

Temporary variations of U.S. households´ foreign equity holdings together with domestic asset wealth are responsible for the adjustment of deviations from the common trend among consumption and aggregate wealth. Foreign equity holdings are predominantly driven by the transitory shock in a cointegrated system consisting of consumption, foreign equity, domestic asset wealth and labour income among which one single cointegration relation exists. This finding suggests that temporary fluctuations of foreign equity holdings must be predictable by the cointegration residual. Evidence presented in this paper reveals that this is indeed the case irrespective if foreign equity holdings are expressed in U.S. dollar terms or in a foreign equity investment weighted compound currency. The predictive power of the cointegration residual is at the same order of magnitude in both cases, which conveys the notion that the adjustment of a transitory deviation from the cointegration trend among consumption and aggregate wealth is only induced by variations of expected returns on the asset wealth components. The correction of a cointegration error through foreign equity holdings is not caused by exchange rate changes but rather by the expectation of returns on foreign equity in the respective local currency and/or changes of the quantity

[12] for evidence of the "equity home bias" see e.g. Tesar and Werner (1995)

of foreign equity that is held by U.S. households. Support for this conclusion is given by long-horizon regressions of excess returns on broad national stock indizes with underlying market capitalization in local currency which mirrors that excess returns, a proxy for equity risk premia, are for most countries better or equally predictable as their counterparts denominated in U.S. dollars which mirrors that expectations of future returns on foreign equity are responsible for the error correction through foreign equity holdings. Exploiting one implication from the error correction mechanism to assess the question if the turnover rate of foreign equity in the U.S. household equity portfolio is higher than the turnover rate of U.S. equity, the conclusion must be drawn that this observation is not mirrored in the U.S. household equity portfolio over the sample period from 1952 to 2004..

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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 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 either directly or indirectly 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 (2001,2004).

- 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 nominal effective exchange rate is constructed as geometrically weighted average of the nominal U.S. dollar spot exchange rates with Australia, Canada, Finland, France, Germany, Hong Kong, Ireland, Italy, Japan, Korea, Mexico, Netherlands, Sweden, Singapore, Spain, Switzerland and United Kingdom. The weights are derived from the IMF´s Coordinated Portfolio Survey of Equity Investment and reflect how large the share of U.S. equity investment in the respective country was in 2001. I assume that this share is constant over the whole sample period. The source of bilateral U.S. dollar spot exchange rates is the

IMF´s International Financial Statistic January 2004. I used the dollar-euro exchange rate for all EMU member countries under consideration since 1999.

- Excess returns on MSCI indizes are defined as real return minus the risk-free rate at the beginning of the period, reflecting the opportunity cost of a U.S. investor investing in foreign equity. Real returns are defined as natural logarithm of the respective index value at the end of period t+1 minus the natural logarithm of the index value at the end of period t. The risk-free rate is the 3-month-U.S. treasury bill. As logarithmic approximations of net returns (continuously compounded returns) are regarded the h-period return is simply the sum of the one period returns over h periods.

Table 1: Johansen Cointegration Test

Critical Values

r =	Trace test 10 %	Trace test 5%	Trace test 1%	L-max test 10 %	L-max test 5%	L-max test 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

Test Statistics

	Trace test	L-max test
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

Information criteria

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

Notes: The variables employed are non-durables and services consumption expenditure 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. All variables are in logarithms, real p.c. in 2000 chain weighted U.S.-dollars.
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: Variance Decompositions (orthogonalized)

$\alpha_c = \alpha_y = 0$

	c_{t+h} - $E_t(c_{t+h})$		fe_{t+h} - $E_t(fe_{t+h})$		daw_{t+h} - $E_t(daw_{t+h})$		y_{t+h} - $E_t(y_{t+h})$	
h	P	T	P	T	P	T	P	T
1	1.0000	0.0000	0.2007	0.7993	0.4169	0.5831	1.0000	0.0000
4	0.9975	0.0025	0.3792	0.6208	0.6019	0.3981	0.9999	0.0001
8	0.9974	0.0026	0.4990	0.5010	0.6683	0.3317	0.9999	0.0001
16	0.9977	0.0023	0.6865	0.3135	0.7466	0.2534	0.9999	0.0001
24	0.9979	0.0021	0.7973	0.2027	0.8022	0.1978	0.9999	0.0001

α_c and α_y estimated

	c_{t+h} - $E_t(c_{t+h})$		fe_{t+h} - $E_t(fe_{t+h})$		daw_{t+h} - $E_t(daw_{t+h})$		y_{t+h} - $E_t(y_{t+h})$	
h	P	T	P	T	P	T	P	T
1	0.9721	0.0279	0.2761	0.7239	0.4719	0.5281	0.9992	0.0008
4	0.9690	0.0310	0.4796	0.5204	0.6735	0.3265	0.9958	0.0042
8	0.9755	0.0245	0.5898	0.4102	0.7388	0.2612	0.9957	0.0043
16	0.9833	0.0167	0.7412	0.2588	0.8044	0.1956	0.9965	0.0035
24	0.9868	0.0132	0.8289	0.1711	0.8463	0.1537	0.9969	0.0031

Notes: This table reports the forecast variance share of the level of the cointegrating variables consumption, c, foreign equity, fw, domestic asset wealth, daw and labour income, y, that can be attributed to the combined three permanent shocks (columns „P") and the single transitory shock (columns „T"). The forecast horizon h is in quarters.

Table 3: Variance Decomposition

	use_{t+h} - $E_t(use_{t+h})$	fe_{t+h} - $E_t(fe_{t+h})$
h=	T	T
1	0.8942	0.7993
4	0.7920	0.6208
8	0.6567	0.5010
16	0.4087	0.3135
24	0.2661	0.2027

Notes: This table reports the fraction of the forecast error variance of U.S. households´ domestic (U.S.) equity holdings, use_{t+h} - $E_t(use_{t+h})$, that can be attributed to the transitory shock in the cointegration relation between consumption and aggregate wealth. The right column presents the forecast error variance decomposition for U.S. households´ foreign equity holdings already reported in table 2 for direct comparison.

Table 4: Forecast Regressions of changes of U.S. households´ foreign equity holdings

h	Δfe_{t+h}	$\Delta neer_{t+h}$	Δfe^{NC}_{t+h}
1	0.9701; R^2: 0.0509 (3.5580)	-0.1189; R^2: -0.0028 (-0.8476)	1.0890; R^2: 0.0542 (3.8472)
4	3.5296; R^2: 0.1986 (3.8210)	-0.6038; R^2: 0.0090 (-0.9717)	4.1334; R^2: 0.2226 (4.4469)
8	6.8687; R^2: 0.3216 (4.7086)	-0.5995; R^2: 0.0017 (-0.8416)	7.4682; R^2: 0.3454 (6.0417)
12	10.4908; R^2: 0.4465 (5.1706)	-0.0606; R^2: -0.0057 (-0.0700)	10.5513; R^2: 0.4335 (5.4026)
14	11.8395; R^2: 0.4505 (5.0127)	0.2833; R^2: -0.0048 (0.2954)	11.5562; R^2: 0.4157 (5.2249)
16	12.7522; R^2: 0.4292 (4.7195)	0.5447; R^2: -0.0024 (0.5100)	12.2076; R^2: 0.3728 (4.8794)
20	14.8859; R^2: 0.4032 (4.5463)	1.1717; R^2: 0.0075 (0.8768)	13.7143; R^2: 0.3108 (4.5842)
24	16.1029; R^2: 0.3608 (4.3291)	2.2608; R^2: 0.0437 (1.1250)	13.8421; R^2: 0.2555 (3.5287)

Notes: Table 2 reports OLS regression results with the cointegration residual as sole regressor. The forecast horizon h is in quarters. Results for forecasts of changes of foreign equity holdings in the U.S. household stock market wealth component in current U.S. dollars, Δfw, are reported in the first column. Changes of the equity investment weighted nominal effective exchange rate, $\Delta neer$, as well as changes of the foreign equity holdings denominated in an equity investment weighted domestic currency basket, Δfw^{NC} are regressed on the cointegration residual as well. Columns two and three display the results. The sample spans the period from second quarter 1952 to second quarter 2004 for Δfw and first quarter 1957 to third quarter 2003 for $\Delta neer$ and Δfw^{NC}. 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 5: Forecast Regressions of excess returns on MSCI stock indizes in US-$

h	1	4	8	12	14	16	20	24
ER_{AUS}	0.8712 (2.1522) R^2: 0.0227	2.8112 (2.3339) R^2: 0.0792	4.0903 (2.9489) R^2: 0.1073	5.4468 (2.7188) R^2: 0.1726	6.1099 (2.7237) R^2: 0.1922	7.1426 (3.1098) R^2: 0.2224	7.4536 (3.0703) R^2: 0.2244	7.3287 (3.1954) R^2: 0.2278
ER_{CND}	0.4319 (1.0964) R^2: 0.0043	1.2131 (0.9291) R^2: 0.0120	2.4105 (1.4618) R^2: 0.0443	4.1117 (3.1971) R^2: 0.1395	4.8221 (3.6636) R^2: 0.1967	5.6523 (4.1470) R^2: 0.2530	6.1120 (3.9067) R^2 0.2181	6.3098 (4.2549) R^2: 0.2551
ER_{FIN}	0.6363 (0.4569) R^2: -0.0056	2.8908 (0.6130) R^2: 0.0121	10.5780 (2.0124) R^2: 0.1199	20.2019 (8.2004) R^2: 0.3059	24.1283 (8.2669) R^2: 0.3760	28.5109 (6.4269) R^2: 0.4152	30.6139 (3.7237) R^2: 0.3529	32.2378 (3.8186) R^2: 0.2161
ER_{FRA}	0.9975 (2.1719) R^2: 0.0303	3.4430 (2.2504) R^2: 0.0919	5.9605 (2.9405) R^2: 0.1421	9.8685 (4.4095) R^2: 0.2982	11.4769 (4.9087) R^2: 0.3523	12.7556 (5.1132) R^2: 0.3699	12.7983 (4.8022) R^2: 0.3345	12.6190 (3.9669) R^2: 0.3166
ER_{GER}	1.0284 (2.6066) R^2: 0.0397	2.8656 (2.3438) R^2: 0.0810	4.9030 (2.0695) R^2: 0.1323	6.8826 (2.1946) R^2: 0.1955	7.2341 (2.1172) R^2: 0.1936	7.4094 (1.8927) R^2: 0.1735	5.6867 (1.4870) R^2: 0.0831	4.7148 (1.3976) R^2: 0.0551
ER_{HK}	0.6945 (0.8630) R^2: -0.0001	2.9609 (1.0041) R^2: 0.0282	5.9392 (1.7642) R^2: 0.0692	7.0278 (2.1702) R^2: 0.0897	7.5809 (2.5882) R^2: 0.1151	7.3780 (2.4916) R^2: 0.1021	5.8164 (1.9486) R^2: 0.0546	4.1227 (1.1030) R^2: 0.0188
ER_{IRL}	1.0663 (2.6858) R^2: 0.0458	3.9169 (4.5234) R^2: 0.2107	7.7882 (4.9633) R^2: 0.4692	9.9650 (3.9062) R^2: 0.4517	11.0006 (3.6189) R^2: 0.4431	13.7069 (4.7835) R^2: 0.4909	15.4923 (7.9547) R^2: 0.3918	12.9429 (6.4231) R^2: 0.1275

Table 5 (continued)

h	1	4	8	12	14	16	20	24
ER_{ITA}	1.0826 (2.1476) R^2: 0.0311	4.2690 (2.3288) R^2: 0.1102	8.5645 (3.0344) R^2: 0.2066	13.4730 (4.1737) R^2: 0.3611	15.6659 (4.6330) R^2: 0.4267	17.7489 (5.4904) R^2: 0.4744	19.7154 (6.8876) R^2: 0.5047	21.0160 (7.5571) R^2: 0.5162
ER_{JPN}	0.3045 (0.5280) R^2: -0.0038	0.8852 (0.3762) R^2: -0.0019	2.7827 (0.7124) R^2: 0.0167	4.5474 (0.9037) R^2: 0.0391	4.9615 (0.9020) R^2: 0.0421	4.3721 (0.7402) R^2: 0.0262	0.6934 (0.1153) R^2: -0.0079	-2.9349 (-0.4645) R^2: 0.0022
ER_{KOR}	-0.3517 (-0.2309) R^2: -0.0147	-4.2011 (-0.8925) R^2: 0.0233	-7.3567 (-1.5213) R^2: 0.0655	-10.6990 (-2.0172) R^2: 0.1534	-12.0323 (-2.2734) R^2: 0.1796	-12.0158 (-2.3204) R^2: 0.1539	-19.6171 (-4.1076) R^2: 0.3274	-33.3904 (-2.8768) R^2: 0.4642
ER_{MEX}	0.6169 (0.6529) R^2: -0.0093	1.4343 (0.5661) R^2: -0.0081	2.5941 (0.6003) R^2: -0.0027	2.7141 (0.5658) R^2: -0.0092	0.6505 (0.1261) R^2: -0.0195	-1.6748 (-0.2460) R^2: -0.0184	-7.4272 (-0.8722) R^2: 0.0089	-20.3085 (-1.3167) R^2: 0.1164
ER_{NL}	1.0005 (3.1157) R^2: 0.0539	3.5965 (3.6089) R^2: 0.2045	6.7214 (4.6815) R^2: 0.3505	9.1531 (4.4616) R^2: 0.4449	9.9173 (4.1096) R^2: 0.4380	10.5030 (3.6327) R^2: 0.4084	9.5943 (3.3981) R^2: 0.2849	8.5472 (3.6851) R^2: 0.2203
ER_{SIN}	0.0440 (0.0579) R^2: -0.0074	-0.2948 (-0.0987) R^2: -0.0071	-0.1769 (-0.0434) R^2: -0.0077	0.4696 (0.1142) R^2: -0.0075	0.8497 (0.2270) R^2: -0.0063	0.9206 (0.2681) R^2: -0.0062	-1.9408 (-0.5987) R^2: -0.0003	-4.5948 (-1.2408) R^2: 0.0354
ER_{ESP}	0.4877 (1.1275) R^2: 0.0015	2.2619 (1.2738) R^2: 0.0305	5.8918 (1.6150) R^2: 0.0879	10.1523 (2.0498) R^2: 0.1513	12.7383 (2.3582) R^2: 0.1979	16.0898 (2.8577) R^2: 0.2620	21.0779 (3.9615) R^2: 0.3514	26.4739 (5.4115) R^2: 0.4582

Table 5 (continued)

h	1	4	8	12	14	16	20	24
ER_{SWE}	0.9822	2.7645	5.4104	9.0320	10.4568	11.6559	12.2607	13.4952
	(1.9047)	(1.6139)	(1.9297)	(2.7881)	(3.1297)	(3.4535)	(4.6541)	(7.2395)
	R^2: 0.0294	R^2: 0.0537	R^2: 0.1347	R^2: 0.2870	R^2: 0.3316	R^2: 0.3563	R^2: 0.3310	R^2: 0.3535
ER_{CH}	0.8574	3.0248	5.2283	7.2651	7.7661	8.1249	6.8636	5.8499
	(2.4347)	(2.6825)	(3.3101)	(3.7095)	(3.3974)	(2.8996)	(2.0392)	(1.8457)
	R^2: 0.0349	R^2: 0.1222	R^2: 0.1767	R^2: 0.2422	R^2: 0.2382	R^2: 0.2190	R^2: 0.1264	R^2: 0.0887
ER_{UK}	1.1559	4.1312	7.5008	10.5090	11.0177	11.8354	10.7710	9.3037
	(3.0031)	(2.8136)	(4.2693)	(6.0194)	(6.1982)	(5.7407)	(4.8302)	(4.1756)
	R^2: 0.0575	R^2: 0.1847	R^2: 0.3510	R^2: 0.4591	R^2: 0.4595	R^2: 0.4659	R^2: 0.3900	R^2: 0.3119

Notes: This table reports OLS estmates for forecast regressions of excess returns on Morgan Stanley Capital International (MSCI) stock indizes with underlying market capitalization in current U.S. dollars. The cointegration residual is the sole regressor. Newey-West corrected t-statistics are displayed in parenthesis. R^2 reports the adjusted R^2 statistic. The forecast horizon h is in quarters. Returns are defined as $r_{t+1} = p_{t+1} - p_t$; where p_t represents the natural logarithm of the respective index value under consideration at the end of period t and p_{t+1} at t+1. Excess returns are defined as $er_t = r_t - r_{f,t}$; with $r_{f,t}$ denoting the risk-free rate at the beginning of period t, here the three-month U.S. treasury bill. As logarithmic approximations are employed the h-period excess return is simply the sum of the excess returns over that time horizon. The sample covers the period from fourth quarter 1969 to second quarter 2004 with the exception of Finland, first quarter 1982 to second quarter 2004, and Ireland, Korea and Mexico, first quarter 1988 to second quarter 2004.

The countries in this sample are Australia, Canada, Finland, France, Germany, Hong Kong, Ireland, Italy, Japan, Korea, Mexico, Netherlands, Singapore, Spain, Sweden, Switzerland and theUnited Kingdom.

Table 6: Forecast Regressions of excess returns on MSCI stock indizes in local currency

h	1	4	8	12	14	16	20	24
ER_{AUS}	0.9266 (2.6657) R^2: 0.0384	3.2216 (2.9836) R^2: 0.1343	5.2512 (3.4087) R^2: 0.2090	7.0189 (3.3911) R^2: 0.2968	7.5932 (3.5216) R^2: 0.3033	8.3696 (3.6840) R^2: 0.3174	8.4029 (3.8692) R^2: 0.2942	7.7399 (3.4903) R^2: 0.2627
ER_{CND}	0.4660 (1.2743) R^2: 0.0091	1.4304 (1.1919) R^2: 0.0247	2.9540 (2.0095) R^2: 0.0783	4.9892 (4.3010) R^2: 0.2000	5.6243 (4.7550) R^2: 0.2434	6.2323 (4.8273) R^2: 0.2804	6.0980 (3.2416) R^2 0.2162	5.4482 (2.3623) R^2: 0.1826
ER_{FIN}	0.3107 (0.2250) R^2: -0.0102	2.0966 (0.4324) R^2: 0.0003	10.3794 (1.7702) R^2: 0.1103	20.8109 (8.1545) R^2: 0.3072	24.5590 (10.0164) R^2: 0.3647	28.3636 (7.6492) R^2: 0.3872	29.8284 (4.0928) R^2: 0.3167	30.1387 (2.9690) R^2: 0.1700
ER_{FRA}	0.8956 (2.3809) R^2: 0.0280	3.2478 (2.4397) R^2: 0.1043	6.2456 (3.6279) R^2: 0.2073	10.5215 (6.2620) R^2: 0.4440	8.7018 (2.7411) R^2: 0.4970	9.1353 (2.5508) R^2: 0.5127	13.4310 (6.8853) R^2: 0.4905	13.2312 (7.2310) R^2: 0.4471
ER_{GER}	0.9928 (3.0440) R^2: 0.0380	2.8814 (2.8861) R^2: 0.0891	5.4218 (2.5626) R^2: 0.1734	8.1692 (2.7923) R^2: 0.2900	9.4555 (4.0941) R^2: 0.3552	9.9331 (4.1620) R^2: 0.2915	8.7733 (2.5125) R^2: 0.2319	9.0559 (3.3135) R^2: 0.2440
ER_{HK}	0.8295 (1.0455) R^2: 0.0038	3.4348 (1.1671) R^2: 0.0429	6.7870 (2.0253) R^2: 0.1002	8.2873 (2.5952) R^2: 0.1446	8.9803 (3.0535) R^2: 0.1856	8.9488 (2.9852) R^2: 0.1724	7.3807 (2.2611) R^2: 0.1053	5.3590 (1.2707) R^2: 0.0439
ER_{IRL}	0.6833 (1.7075) R^2: 0.0036	2.8015 (2.7008) R^2: 0.0772	7.8372 (4.4004) R^2: 0.3672	11.8549 (6.0859) R^2: 0.5256	13.1840 (5.4869) R^2: 0.5384	16.1804 (6.6709) R^2: 0.5910	19.3477 (11.1509) R^2: 0.5238	20.2432 (7.2006) R^2: 0.2781

Table 6 (continued)

h	1	4	8	12	14	16	20	24
ER_{ITA}	0.9426	3.8837	8.2274	12.7819	14.5697	16.0367	16.9051	16.8953
	(2.1360)	(2.3985)	(3.6691)	(6.1464)	(7.1690)	(8.2177)	(7.9798)	(5.8199)
	R^2: 0.0240	R^2: 0.1033	R^2: 0.2252	R^2: 0.3927	R^2: 0.4391	R^2: 0.4592	R^2: 0.4151	R^2: 0.3515
ER_{JPN}	0.1989	0.4271	1.5949	2.6973	2.9343	2.7769	0.8402	-1.9108
	(0.4351)	(0.2448)	(0.5878)	(0.7836)	(0.7745)	(0.6770)	(0.2108)	(-0.4799)
	R^2: -0.0052	R^2: -0.0053	R^2: 0.0078	R^2: 0.0246	R^2: 0.0257	R^2: 0.0177	R^2: -0.0068	R^2: -0.0012
ER_{KOR}	-0.2240	-3.3080	-5.0060	-6.6957	-7.5309	-6.6898	-11.7434	-21.1220
	(-0.1767)	(-0.8771)	(-1.4601)	(-1.8554)	(-2.1409)	(-1.9527)	(-3.5109)	(-2.3842)
	R^2: -0.0151	R^2: 0.0239	R^2: 0.0491	R^2: 0.0920	R^2: 0.1080	R^2: 0.0750	R^2: 0.2386	R^2: 0.3746
ER_{MEX}	1.3986	4.1022	7.1126	8.6184	7.5555	7.3629	3.4299	-11.8020
	(2.1765)	(2.4925)	(2.5254)	(2.5154)	(1.9624)	(1.4567)	(0.4938)	(-0.7587)
	R^2: 0.0381	R^2: 0.0956	R^2: 0.1478	R^2: 0.1154	R^2: 0.0676	R^2: 0.0426	R^2: -0.0143	R^2: 0.0209
ER_{NL}	0.9602	3.5831	7.1335	10.2756	11.1911	11.9504	12.2255	12.1612
	(3.3022)	(3.6431)	(4.8319)	(4.6533)	(4.4992)	(4.1560)	(4.0722)	(4.5859)
	R^2: 0.0505	R^2: 0.1847	R^2: 0.3484	R^2: 0.4864	R^2: 0.4884	R^2: 0.4813	R^2: 0.4218	R^2: 0.3819
ER_{SIN}	0.0736 -	-0.1427	0.2467	1.0871	1.5117	1.7224	-0.4926	-2.6078
	(0.1025)	(-0.0508)	(0.0680)	(0.3220)	(0.5178)	(0.6636)	(-0.2022)	(-0.9897)
	R^2: -0.0073	R^2: -0.0074	R^2: -0.0077	R^2: -0.0039	R^2: 0.0003	R^2: 0.0024	R^2: -0.0078	R^2: 0.0143
ER_{ESP}	0.5208	2.4722	6.7432	11.5497	13.7213	16.4466	20.6856	25.0703
	(1.4210)	(1.7083)	(2.3969)	(3.1202)	(3.4419)	(3.9288)	(5.4473)	(7.7154)
	R^2: 0.0036	R^2: 0.0501	R^2: 0.1676	R^2: 0.2828	R^2: 0.3226	R^2: 0.3785	R^2: 0.4514	R^2: 0.5383

Table 6 (continued)

h	1	4	8	12	14	16	20	24
ER_{SWE}	0.9363	2.7052	6.0884	10.4780	11.9249	13.0336	13.7147	14.6489
	(2.1320)	(1.9210)	(2.5817)	(3.7663)	(4.2812)	(4.7144)	(7.4838)	(7.9685)
	R^2: 0.0235	R^2: 0.0482	R^2: 0.1476	R^2: 0.3115	R^2: 0.3460	R^2: 0.3650	R^2: 0.3470	R^2: 0.3457
ER_{CH}	0.8786	3.3948	6.5848	9.6324	10.4105	11.5120	12.3286	13.1869
	(2.9600)	(3.4649)	(4.7978)	(5.4072)	(4.9490)	(4.6089)	(4.0036)	(4.1445)
	R^2: 0.0408	R^2: 0.1759	R^2: 0.3220	R^2: 0.4723	R^2: 0.4665	R^2: 0.4754	R^2: 0.4069	R^2: 0.3984
ER_{UK}	1.0514	3.6953	6.6144	8.8991	8.9403	9.2532	7.5599	5.8112
	(3.0164)	(2.6453)	(4.1872)	(5.2282)	(4.7869)	(4.1870)	(4.0982)	(4.3172)
	R^2: 0.0544	R^2: 0.1703	R^2: 0.3271	R^2: 0.4269	R^2: 0.4075	R^2: 0.4057	R^2: 0.2866	R^2: 0.1789

Notes: This table reports OLS estmates for forecast regressions of excess returns on Morgan Stanley Capital International (MSCI) stock indizes with underlying market capitalization in current local currency. The cointegration residual is the sole regressor. Newey-West corrected t-statistics are displayed in parenthesis. R^2 reports the adjusted R^2 statistic. The forecast horizon h is in quarters. Returns are defined as $r_{t+1} = p_{t+1} – p_t$; where p_t represents the natural logarithm of the respective index value under consideration at the end of period t and p_{t+1} at t+1. Excess returns are defined as $er_t= r_t– r_{f,t}$; with $r_{f,t}$ denoting the risk-free rate at the beginning of period t, here the three-month U.S. treasury bill. As logarithmic approximations are employed the h-period excess return is simply the sum of the excess returns over that time horizon. The sample covers the period from fourth quarter 1969 to second quarter 2004 with the exception of Finland, first quarter 1982 to second quarter 2004, and Ireland, Korea and Mexico, first quarter 1988 to second quarter 2004.

The countries in this sample are Australia, Canada, Finland, France, Germany, Hong Kong, Ireland, Italy, Japan, Korea, Mexico, Netherlands, Singapore, Spain, Sweden, Switzerland and theUnited Kingdom.