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

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Abstract

A simple manipulation of the cointegrated framework proposed by Lettau and Ludvigson (2001, 2004) allows to demonstrate that temporary fluctuations of the U.S. consumption-wealth ratio predict excess returns on international stock markets. This finding is the reflection of an important common, temporary component in international stock markets and thus provides empirical evidence for a robust link between stock markets at business cycle frequency. Moreover, I find that between one third and more than a half of the covariation of long-horizon returns on the G7 stock markets is explained by the common transitory stock market component identified in this paper. Furthermore, U.S. households seem to rebalance their foreign equity portfolio in response to the perception of local currency rather than exchange rate adjusted returns.

1 Introduction

Long-term predictability of asset returns is by now well documented in a growing body of empirical literature.[1] This paper contributes to this literature by employing the framework proposed by Lettau and Ludvigson (2001, 2004) to explore the cyclical link between international stock markets.

Lettau and Ludvigson (2001, 2004) show that fluctuations of the U.S. consumption-wealth ratio, the residual of the cointegration relation between consumption and aggregate wealth, predict real and excess returns on broad U.S. stock indexes. Using the Lettau and Ludvigson framework, Fernandez-Corugedo et al. (2003), Fisher and Voss (2004) and Tan and Voss (2003) provide evidence for the forecast ability of the UK and Australian consumption-wealth ratio for excess returns on respective national stock indexes. Hamburg et al. (2005) find that German stock market returns are not predicted by variations of the national consumption-wealth ratio which seems to be the consequence of limited stock market participation of German households.

Apart from the special case of Germany, consumption-wealth ratios of Anglo-Saxon countries predict excess returns on national stock markets and thus corroborate theoretical macroeconomic models which argue that time variation of stock returns is due to cyclically varying risk premia. Fluctuations of risk premia over time result from agents´ time-varying risk aversion over the business cycle induced by the formation of consumption habits (Campbell and Cochrane, 1999) or the presence of uninsurable background risks (Constantinides and Duffie, 1996; Heaton and Lucas, 2000 a, b).

But despite macroeconomic explanations of stock return predictability and evidence of strong comovement between national stock markets since the mid-eighties (Brooks and Del Negro, 2005), a robust link between stock markets at business cycle frequency seems hard to capture. Furthermore, it is not even clear how comovement of stock markets can be rationalized. Do international stock markets follow a common stochastic trend as argued by Kasa (1992) or is comovement rather the outcome of a common stationary component in stock prices? Richards (1995) weakens the basis of Kasa´s statistical evidence which gives place for scepticism regarding evidence of a common stochastic trend among international stock markets. However, so far the literature also lacks convincing evidence of a common stationary component in stock markets.

[1] see Cochrane (2005) for an excellent survey

In this paper I argue that the cointegration framework proposed by Lettau and Ludvigson (2001, 2004) has the potential to shed light on this issue. Based on the idea that transitory fluctuations of wealth leave consumption unaffected, Lettau and Ludvigson provide evidence that mainly transitory market value changes of U.S. households´ stock holdings cause the U.S. consumption-wealth ratio to fluctuate temporarily. These market value changes are induced by the expectation of time-varying stock returns which explains the predictive power of short-run variations of the U.S. consumption-wealth ratio for excess returns on the U.S. stock market. U.S. households´ stock market wealth is a prime example of the home bias in equity portfolios (Tesar and Werner, 1995). Nevertheless, U.S. households hold either directly or indirectly foreign stocks which amounts to a relatively small part of U.S. households´ stock market wealth, but inevitably raises the question if the explanation for the predictive power of the U.S. consumption-wealth ratio for the U.S. stock market also pertains to foreign stock markets.

I deal with this issue by using a simple manipulation of the Lettau and Ludvigson framework to show that variation in the market value of U.S. households´ foreign equity holdings is induced by the expectation of time-varying returns on foreign stock markets. Hence, temporary variations of the U.S. consumption-wealth ratio predict excess returns on international stock markets. These findings leave the impression that the ratio of consumption to aggregate wealth in the U.S. reflects a common, temporary component in international stock markets. I present evidence that between one third and more than a half of the covariation of long-horizon returns on G7 stock indexes can be attributed to the common, transitory stock market component.

Furthermore, exchange rate changes are not predictable by variations of the U.S. consumption-wealth ratio which conveys the notion that exchange rate changes do not cause the cyclical market value variation in U.S. households' foreign equity holdings. This finding underlines that the U.S. consumption-wealth ratio echoes a common stock market component.

Additionally, I provide evidence that U.S. households rebalance their equity portfolio in response to the perception of time variation in expected local currency returns rather than to exchange rate adjusted returns.

The remainder of this paper is organised as follows. The theoretical framework is introduced in section two. Section three discusses the cointegration and error correction properties of my U.S. consumption-wealth ratio approximation in detail. Section four identifies permanent and transitory shocks in the cointegrated system and reports variance decompositions of the cointegrated variables with respect to these shocks. Section five provides details about long-horizon regressions of changes of U.S. households´ foreign equity holdings and excess returns on foreign stock indexes on the cointegration residual. Section six gives evidence for the robustness of the long-horizon regressions and quantifies to what extent the common stock market component is responsible for the covariation of long-horizon returns on G7 stock markets. Section seven concludes. The appendix contains a detailed description of data employed in this paper.

2 The Consumption-Wealth Ratio

I follow Lettau and Ludvigson (2001,2004) as well as Campbell and Mankiw (1989) and consider a representative agent economy in which all wealth is traded. The representative household faces an intertemporal budget constraint of the form

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

where W_t denotes aggregate wealth (human wealth plus asset wealth) in period t, C_t, consumption and $R_{w,t+1}$ the net return on aggregate wealth.

Rearranging the budget constraint for the ratio of consumption to wealth and taking a loglinear approximation around the mean consumption-wealth ratio under the assumption that this mean is covariance stationary leads to the following law of motion for the log consumption-wealth ratio.

$$c_t - w_t = E_t \sum_{i=1}^{\infty} \rho_w^i (r_{w,t+i} - \Delta c_{t+i}) \tag{2}$$

Lower-case letters denote natural logarithms throughout the paper, Δ represents the difference operator. In order to make (2) empirically tractable, Lettau and Ludvigson (2001, 2004) decompose aggregate wealth into its components asset and human wealth and loglinearise around the long-run mean of the ratio of human and asset wealth which leads to

$$w_t \approx v a_t + (1 - v) h_t \tag{3}$$

with v interpretable as average share of asset wealth in aggregate wealth, a_t, log asset wealth and h_t, log human wealth. I further decompose asset wealth into foreign equity held by U.S. households and the rest of asset wealth which

I will refer to as domestic asset wealth. A loglinear approximation of asset wealth around the foreign equity to domestic asset wealth ratio yields

$$a_t \approx \lambda fe_t + (1-\lambda)daw_t \tag{4}$$

with λ the average share of foreign equity in U.S. households´ asset wealth, fe_t, foreign equity and daw_t, domestic asset wealth.

Combining (3) with (4) gives

$$w_t \approx \theta fe_t + \phi daw_t + (1-\theta-\phi)h_t \tag{5}$$

where $\theta = v\lambda$ is the average share of foreign equity in aggregate wealth and $\phi = v(1-\lambda)$ the average share of domestic asset wealth in aggregate wealth.

A loglinear approximation of the gross return on asset wealth with respect to foreign equity and domestic asset wealth combined with the loglinear proxy of the return on aggregate wealth decomposed into asset wealth and human wealth[2] gives

$$r_{w,t} = \theta r_{fe,t} + \phi r_{daw,t} + (1-\theta-\phi)r_{h,t} \tag{6}$$

Plugging (6) and (4) into (2) and taking expectations on both sides of the equation yields

$$\begin{aligned} & c_t - \theta fe_t - \phi daw_t - (1-\theta-\phi)h_t \\ = \; & E_t\{\sum_{i=1}^{\infty}\rho_w^i[(\theta r_{fe,t+i} + \phi r_{daw,t+i} + (1-\theta-\phi)r_{h,t+i}) - \Delta c_{t+i}]\} \end{aligned} \tag{7}$$

However, (7) cannot be employed for empirical purposes because one part of aggregate wealth, human wealth, is unobservable. I assume labour income to represent the dividend paid from human wealth and thus its non-stationary component to overcome this obstacle. Then log human wealth, h_t, obeys

$$h_t = \kappa + y_t + z_t \tag{8}$$

with, y_t, log labour income, κ, a constant term and a covariance stationary term z_t. Plugging (8) into (7) and assuming that the net return on labour income equals the net return on human wealth gives

[2] see Campbell (1996) for details

$$
\begin{aligned}
& c_t - \theta fe_t - \phi daw_t - (1-\theta-\phi) y_t && (9) \\
= \quad & E_t \{ \sum_{i=1}^{\infty} \rho_w^i [(\theta r_{fe,t+i} + \phi r_{daw,t+i} + (1-\theta-\phi)\Delta y_{t+i}) - \Delta c_{t+i}] + (1-\theta-\phi) z_{t+i} \}
\end{aligned}
$$

According to (9) c_t, log consumption, fe_t, log foreign equity, daw_t, log domestic asset wealth and y_t, log labour income cointegrate, provided they are integrated of order one, I(1). Hence, time variation of the consumption-wealth ratio, i.e. a temporary deviation from the common trends, mirrors either changes of (returns on) foreign equity, returns on domestic asset wealth, changes of labour income or consumption growth, or an arbitrary combination. Furthermore, estimates of the cointegration coefficients should reflect θ, ϕ and $(1-\theta-\phi)$, the average shares of the wealth components in total wealth.

2.1 Empirical evidence: Cointegration and error correction

In this section, I assess the cointegration properties of my loglinear proxy of the U.S. consumption-wealth ratio. All variables are quarterly, per capita, real in billions of chain-weighted 2000 U.S. dollars and transformed to natural logarithms for the sample period from second quarter 1952 to second quarter 2004.

As pointed out by Lettau and Ludvigson (2001) as well as Rudd and Whelan (2002), the budget constraint (1) refers to total personal consumption flows. Since we do not observe consumption flows we rely on expenditures as best proxy. I thus follow Blinder and Deaton (1985) and approximate log total consumption expenditure as constant multiple of log non-durables and services consumption expenditure excluding clothing and shoes.

Rudd and Whelan (2002) provide evidence that the linear relation between log non-durable and services consumption and log total personal consumption expenditure is not constant over time. However, Lettau and Ludvigson (2004) argue that durable consumption expenditure represents rather replacements or additions to an existing stock than a service flow from the stock of durable goods and hence is better described as wealth which is the view I follow in this paper. Furthermore, Rudd and Whelan cast doubt on the appropriateness of using different deflators to obtain real variables. Lettau and Ludvigson use the deflator for total personal consumption expenditure to deflate their asset wealth and labour income proxy but a different deflator for their consumption approximation. I take this critique into account and deflate all variables with the CPI of total personal consumption expenditure.

Labour income is proxied as proposed by Lettau and Ludvigson (2001, 2004). U.S. households´ foreign equity holdings are determined as explained in detail in the appendix. Domestic asset wealth is household net worth less foreign equity holdings.

Unit root tests provide evidence that each variable employed in this analysis contains a unit root. In addition, I cannot reject that first differences of the variables under consideration are stationary.[3] Non-durable consumption, foreign equity, domestic asset wealth and labour income are I(1), which conveys the notion that my four-variable approximation of the log consumption-wealth ratio should cointegrate. Table 1 displays results of the Johansen cointegration test, critical values for Trace and L-max test as well as the test statistics for both of the tests. Akaike (AIC) and Schwartz (SIC) information criteria suggest an appropriate lag length of one quarter for the vector autoregressive representation (VAR) of the four variables. According to the test statistics, I cannot reject the null of no cointegration for the relation between non-durables and services consumption expenditure excluding clothing and shoes, foreign equity holdings, domestic asset wealth and labour income at 90 percent confidence level.[4]

However, Hoffmann and Mac Donald (2003) point out that the existence of a cointegrating relationship cannot be only grounded on statistical devices but should incorporate economic theory. Furthermore, theory suggests that the estimates of the cointegration coefficients should reflect the average share of the respective wealth component in total wealth. I impose one cointegrating relationship on consumption, foreign equity, domestic asset wealth and labour income and estimate the cointegration vector to investigate this point.

As emphasized by Stock (1987), 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. But the error terms of the individual time-series variables could be correlated with each other. Hence,

[3] Results are not reported but available upon request.

[4] Philipps-Ouliaris cointegration test as well as the cointegration test by Shin provide qualitatively similar results which are not reported but available from the author upon request.

the OLS estimates are consistent but potentially biased away from the true values.

That is why I follow Stock and Watson (1993) who propose a dynamic least squares technique to overcome this obstacle which is in this context achieved by adding leads and lags of first differences of foreign equity, domestic asset wealth and labour income as additional regressors in a regression of consumption on the level of the three wealth components. The estimate equation takes the following form:

$$\begin{aligned} c_t &= \alpha + \beta_{fe} fe_t + \beta_{daw} daw_t + \beta_y y_t \\ &\quad + \sum_{i=-k}^{k} b_{fe,i} \Delta fe_{t-i} + \sum_{i=-k}^{k} b_{daw,i} \Delta daw_{t-i} + \sum_{i=-k}^{k} b_{y,i} \Delta y_{t-i} + \varepsilon_t \end{aligned} \tag{10}$$

Estimation of the cointegration coefficients β_i with $i = fe, daw, y$ gives $\widehat{\boldsymbol{\beta}}^{ndc}$ if the coefficient on non-durable consumption is normalized to unity with t-statistics in parenthesis.[5] The coefficients of differences in lead or lag are omitted.

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

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, Hoffmann (2005) provides an explanation for this finding. Equation (10) is derived from the budget constraint (1) which refers to total consumption. I assume total personal consumption expenditure to be a constant multiple of non-durables and services consumption, i.e. total personal consumption less consumption expenditure of durable goods on the left hand side of (10). But the stock of durables is included in the asset wealth proxy on the right hand side of the estimate equation such that the estimates should sum to a number larger than one. Estimation of the cointegration vector using total personal consumption instead of non-durables and services consumption expenditure

[5]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.

leads to the cointegration vector $\widehat{\boldsymbol{\beta}}^{tc}$.

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

Note that the cointegration coefficient estimates of foreign equity and labour income remain relatively stable whereas the coefficient of domestic asset wealth less the stock of durable goods decreases. The sum of the wealth cointegration coefficients is now almost exactly unity.

If non-durable and services consumption expenditure is used as consumption proxy, then the sum of the wealth cointegration coefficients increases by eight percent to 1.08. An interpretation of this finding is that the present value of durable consumption amounts to eight percent of the present value of total consumption. Hence, estimates of the cointegration vector employing non-durable consumption expenditures should mirror that in the long-run the net present value of asset wealth including the stock of durable goods exceeds the present value of non-durable consumption by approximately eight percentage points. This argument is consistent with the respective cointegration coefficient estimates of asset wealth inclusive and exclusive the stock of durable goods.

Furthermore, the Federal Reserve Board of Governors reports replacement costs of durable goods in household net worth. The share of durable goods in household net worth is around eight to nine percent on average over the sample period which further supports the reasoning from above.

The point estimate of the foreign equity cointegration coefficient is reasonable as well, i.e., it mirrors the average share of foreign equity in total wealth over the sample period from 1952 to 2004.

Based on the economically meaningful cointegration coefficient estimates, I assume the presence of one cointegrating relationship between the four variables under consideration throughout the paper.

I thus proceed to assess the error correction properties of the cointegrated system with non-durable and services consumption expenditure as consumption proxy to examine if foreign equity adjusts a temporary deviation from the common trend among consumption and aggregate wealth. I exploit that for every cointegrating 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

$$\mathbf{\Gamma(L)\Delta x}_t = \boldsymbol{\alpha}\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{11}$$

in which $\mathbf{\Delta x}_t = (\Delta c_t, \Delta fe_t, \Delta daw_t, \Delta y_t)'$ is the vector of first differences and $\mathbf{x}_{t-1}$ the vector of lagged levels, $\boldsymbol{\alpha} = (\alpha_c, \alpha_{fe}, \alpha_{daw}, \alpha_y)'$ represents the vector of error correction coefficients. $\mathbf{\Gamma(L)}$ denotes a (4 by 4) matrix polynomial in the lag operator and $\widehat{\boldsymbol{\beta}} = (1, -\widehat{\beta}_{fe}, -\widehat{\beta}_{daw}, -\widehat{\beta}_y)'$ is the vector of the above estimated cointegration coefficients when non-durable and services consumption expenditure is used as consumption proxy. Hats indicate estimated variables and $\boldsymbol{\varepsilon}_t$ represents the (4 by 1) 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. If $\mathbf{x}_t$ is cointegrated, at least one of the adjustment coefficients $\alpha_c, \alpha_{fe}, \alpha_{daw}$ or α_y must take values different from zero in the error-correction representation.

Table 2 reports VECM coefficient estimates. The lag length of one has been chosen according to Akaike and Schwartz information criteria. T-statistics of the coefficient estimates are in parenthesis. I focus on the adjustment coefficients in the last row of table 2.

The adjustment coefficient estimates of both asset wealth components are statistically different from zero which mirrors the responsibility of asset wealth for the error correction in the cointegrated system. Domestic asset wealth adjusts temporary deviations from the common trend between consumption and total wealth which is presumably driven by the domestic stock market wealth component (Lettau and Ludvigson, 2004).

In addition, the adjustment coefficient of foreign equity is not only statistically significant but also relatively high. One might be concerned about that estimate which implies a fast correction of the cointegration error through foreign equity and seems to be too high compared to the adjustment coefficient of domestic asset wealth or the error correction coefficient of total asset wealth in Lettau and Ludvigson (2001, 2004). However, foreign equity, one component that is particularly responsible for the adjustment of a temporary deviation from the common trends, is isolated from the rest of asset wealth. Lettau and Ludvigson (2004) identify the stock market component of asset wealth to be predominantly driven by the transitory shock, whereas the permanent shocks are to most extent responsible for variations of non-stock market wealth. Hence, the adjustment coefficient of stock market and non-stock market wealth combined should be substantially lower than that of an isolated stock market wealth component. Moreover, data on household net worth published by the Federal Board of Governors discloses that on average non-stock market wealth accounts for 78 percent of asset wealth over the sample period from second quarter 1952 to second quarter 2004. A high adjustment coefficient of foreign equity compared to domestic asset wealth is hence reasonable since domestic asset wealth is dominated by non-stock market wealth.

The negative signs of the consumption and labour income coefficients are not particularly worrisome as well because they are not statistically distinguishable from zero.

2.2 Identification of permanent and transitory shocks and variance decomposition

I follow Hoffmann (2001) in identifying permanent and transitory shocks in the cointegrating system in order to quantify their contribution to the forecast error variance of the level of consumption, foreign equity, domestic asset wealth and labour income and to give further evidence for the robustness of the results from the previous section.

As I regard a cointegrated system with four variables and one single cointegrating 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

$$\mathbf{x}_t = \mathbf{C(1)} \sum_{i=0}^{t} \varepsilon_i + \mathbf{C}^*(\mathbf{L}) \boldsymbol{\varepsilon}_t \tag{12}$$

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

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

$$\mathbf{C}(\mathbf{1}) = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma}(\mathbf{1})\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp} \tag{13}$$

in which $\boldsymbol{\beta}_{\perp}, \boldsymbol{\alpha}_{\perp}$ 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 $\mathbf{C}(\mathbf{1})\boldsymbol{\alpha} = 0$. The common trends, π_t, thus are

$$\pi_t = \boldsymbol{\alpha}'_{\perp}\sum_{i=0}^{t}\varepsilon_i = \sum \eta_t. \tag{14}$$

Let $\eta_t^P = \boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ 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 are identified via

$$\boldsymbol{\eta}_t = \mathbf{S}\boldsymbol{\varepsilon}_t \tag{15}$$

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 three permanent shocks and the single transitory shock to the forecast error variance of the four cointegrated variables.

Table 3 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. I identify the transitory shock as orthogonal to the three permanent shocks. The top panel reports the variance decomposition if statistically insignificant adjustment coefficient estimates are set to zero as recommended by Gonzalo and Ng (2001). The bottom panel displays the variance decomposition if all adjustment coefficients are set to their estimated values.

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 of the variables 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. 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 in line with the magnitude of the error correction 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. Consumption and labour income do not participate in the error correction. Their adjustment coefficients are statistically indistinguishable from zero, which means that both variables should be predominantly driven by the permanent shocks. Variance decompositions for consumption and labour income support this reasoning. Almost all of the variation of consumption and labour income can be attributed to the three permanent shocks at any time horizon.

3 Forecasting power of the cointegration residual

The estimated adjustment coefficients and the variance decompositions imply that the cointegration residual should serve as a predictor of changes of U.S. households´ foreign equity holdings.[6] I perform long-horizon regressions with the cointegration residual as sole regressor to assess this point. The long-horizon regressions take the following general form

$$\sum_{h=1}^{H} \Delta y_{t+h} = \mu + \beta_{t+h} x_t + \varepsilon_{t+h} \tag{16}$$

where x is the cointegration residual, y the natural logarithm of the regressand, μ denotes a constant and ε the error terms at the respective time horizon $t+h$.

I focus on in-sample regressions to provide evidence for predictability throughout this paper since out-of sample regressions do not necessarily provide superior, more robust, results in favour or against predictability (Inoue and Kilian, 2004). But before describing the evidence it may be useful to provide some intuition of what should be reflected in the regression outcomes. A temporarily high consumption-wealth ratio is associated with the expectation of high future returns on aggregate wealth. The positive sign of the foreign equity error correction coefficient estimate suggests that we should expect positive regressor estimates in forecast regressions of market value changes of foreign equity holdings on the cointegration residual because

[6] Lettau and Ludvigson (2004) show that the cointegration residual neither predicts consumption nor labour income growth.

a temporarily high U.S. consumption-wealth ratio is associated with a high market value of foreign equity holdings.

High expected returns on foreign stock markets induce households to increase their foreign equity investment and thus the market value of their foreign equity holdings. Hence, positive regressor estimates in forecast regressions of excess returns on foreign stock indexes on the cointegration residual would be the consequence. However, this reasoning does not necessarily have to apply to all foreign stock markets. So, I cannot preclude negative estimates in regressions of national stock returns on the cointegration residual.

The left column of table 4 presents estimates from the regression of changes of U.S. households´ foreign equity holdings, $\Delta fe^{\$}$, on $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ with Newey-West corrected t-statistics in parenthesis. The forecast horizon, h, is in quarters. All regressor coefficient estimates are statistically significant. The R^2 statistic peaks at 14 quarters and displays that the cointegration residual explains 45 percent of the variation of foreign equity holdings in U.S. wealth.

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 local currency or changes of the nominal U.S exchange rate vis-à-vis the rest of the world or an arbitrary combination are responsible for a correction of the cointegration error and hence predictable. In order to shed light on this issue, I construct a foreign equity investment weighted effective exchange rate of the U.S. dollar relative to the countries the U.S. hold equity of. I focus on countries in which the U.S. invest at least one percent of their foreign equity investment.[7] The 17 countries used in this analysis represent about 80% of U.S. foreign equity investment. Data on foreign equity investment is from the IMF´s coordinated portfolio survey 2001. I use the share of U.S. equity investment into a particular country from total U.S. foreign equity investment as a weight to construct the effective exchange rate. I assume these weights, derived from 2001 data, to be constant over the sample period from first quarter 1957 to third quarter 2003. This assumption certainly biases the effective exchange rate towards the foreign equity investment pattern of the U.S. in

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

recent years. However, foreign equity flows have grown substantially since the last twenty years (Hau and Rey (2004)), such that shares of U.S. foreign equity investment in recent years appropriately display where and to what extent the U.S. invest in foreign equity. The weights described above should thus be sufficient to approximate the true foreign equity investment weighted effective U.S. dollar exchange rate.

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 the cointegration residual will forecast changes of the equity investment weighted exchange rate. Moreover, I investigate if changes of foreign equity holdings denominated in a weighted basket of national currencies are predictable. I employ the equity investment weighted exchange rate to obtain foreign equity holdings in such a compound currency. Predictability of changes of these holdings either reflects variation of equity prices in local currency or changes of the quantity of foreign equity holdings or both. However, these effects cannot be distinguished in this exercise.

The middle column of table 4 provides evidence that changes of the effective exchange rate, $\Delta neer$ are not predicted by the cointegration residual for the time period from first quarter 1957 to third quarter 2003. None of the regressor coefficients is statistically distinguishable from zero. This result is corroborated by not reported regressions of bilateral exchange rate changes on the cointegration residual.[8]

The non-predictability of exchange rate changes leaves the impression that $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ predicts changes of foreign equity holdings in local currency, Δfe^{NC}. The right column of table 4 presents the regression results. All regressor coefficients are statistically distinguishable from zero. The peak of predictability is reached after 12 quarters explaining 43 percent of foreign equity holdings variation if they are expressed in a weighted basket of national currencies. The evidence of predictability is as strong as in the case of foreign equity holdings denominated in U.S. dollar.

The forecast regressions reported in table 4 reveal that the cointegration residual displays information about market value changes of U.S. households´ foreign equity holdings. The non-predictability of exchange rate changes conveys the notion that the correction of the cointegration error through foreign equity holdings is caused by the expectation of time-varying returns on foreign stock markets and subsequent portfolio rebalancing. Exchange

[8]Results are available upon request.

rate changes play a negligible role in the error correction mechanism.

At this stage, however, it is not clear whether households react to the perception of exchange rate adjusted expected returns or to returns in local currency. In order to assess this point I examine if $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ predicts excess returns on Morgan Stanley Capital International (MSCI) stock indexes for the countries employed to calculate the effective exchange rate. The MSCI indexes offer the advantage that the methodology of index construction is the same for all country indexes and hence allows for direct comparison. Furthermore, MSCI publishes index values with underlying market capitalisation denominated in U.S. dollar and local currency.

Table 5 reports estimates from long-horizon regressions when excess returns on MSCI indexes with underlying market capitalisation in U.S. dollars are regressed on the cointegration residual. Newey-West corrected t-statistics appear in parenthesis. The forecast horizon, h, is in quarters. The sample covers the period from 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. Excess returns are defined as simple return on a country index less the three-month U.S. treasury bill rate at the beginning of period, reflecting the U.S. investor´s opportunity cost of investing in a foreign stock market.

The regression results of table 5 mirror that the cointegration residual predicts excess returns with underlying market capitalisation in U.S. dollars best at 8 to 24 quarter frequency. With the exception of Japanese, Mexican and Singaporean stock index returns, which are not predictable at any time horizon, 14 MSCI stock index returns in U.S. dollars are predictable. The predictive power of the cointegration residual differs widely across countries, at most 11,5 percent of the variation of excess returns on the Hong Kong MSCI index is explained by $\widehat{\boldsymbol{\beta}}'\mathbf{x}_{t-1}$ compared to 51,6 percent for Italy. A notable outlier is Korea, which displays forecastability at long horizons, 12 to 24 quarters, but the regressor coefficient is negative, in contrast to the economic intution given at the beginning of the section. However, an explanation for this finding is straightforward. During the short sample period for which data on the Korean MSCI stock index is available, 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 estimates. A high U.S. consumption-wealth ratio is thus associated with negative returns on the Korean stock market.

Noteworthy as well is the predictive power of $\widehat{\boldsymbol{\beta}}' \mathbf{x}_{t-1}$ for excess returns on the German MSCI index. It explains about 20% of the German stock index return variation. The U.S. consumption-wealth ratio reflects fluctuations of the German stock market, whereas the German consumption-wealth ratio displays fluctuations of the German unemplyoment rate and other business cycle variables (Hamburg et al. (2005)).

Table 6 reports estimates from forecast regressions of excess returns on MSCI stock indexes with underlying market capitalisation in local currency on the cointegration residual. The overall picture that emerges is that excess returns in national currency are at least equally predictable as returns denominated in U.S dollars. The predictive power of the cointegration residual for excess returns in local currency leaves the impression that temporary fluctuations of foreign equity holdings are induced by the expectation of expected returns in local currencies rather than exchange rate adjusted returns.

4 Robustness check and implications for stock market comovement

The long-horizon regressions in the previous section suggest that short-run fluctuations of the U.S. consumption-wealth ratio explain close to one half of the variation in long-horizon returns on some of the foreign country indexes. This result suggests that $\widehat{\boldsymbol{\beta}}' \mathbf{x}_{t-1}$ reflects an important common, cyclical component in international stock markets which could be interpreted as risk factor common to international stock market returns. But how reliable are the results from the forecast regressions? Does variation in the U.S. consumption-wealth ratio really capture short-run movements of international stock markets? Or are the high R^2 statistics obtained in the long-horizon regressions spurious?

Figures 1 (a) to (h) plot actual realisations of 16-quarter returns on country indexes for which the forecast regressions provided R^2 statistics of around 0.4 or higher together with the fitted values of the cointegration residual for that time horizon. Note that the return series are denominated in local currency. The countries in question are France, Ireland, Italy, Netherlands, Spain, Switzerland, United Kingdom and the United States. I consider returns from beginning of the 1970s to present for all countries except Ireland.

The figures show that the high R^2 statistics for long-horizon returns on the country indexes under consideration indeed reflect the explanatory power of variation in the U.S. consumption-wealth ratio for short-term movements of international stock markets. The relationship between actual long-horizon returns and fitted values of the cointegration residual is far from being spurious. Quite in contrast, the fitted short-run fluctuations of the ratio between consumption and aggregate wealth in the U.S. provide a good description of the variation in returns on international stock markets. Hence, the figures support the view that the U.S. consumption-wealth ratio echoes a common, transitory component in international stock markets.

Thus the cointegration residual seems to represent a risk factor common to international stock markets which explains a considerable fraction of the variation in long-horizon returns. As this risk factor is common to stock markets and mirrored in the U.S. consumption-wealth ratio, we could obtain information about the degree of covariation between international stock markets from

$$var(\mathbf{r}) = \gamma\gamma' var(cay) + cov(\varepsilon) \tag{17}$$

where $\mathbf{r}$ is the vector of long-horizon returns, cay represents the cointegration residual, i.e. the common risk factor and $\boldsymbol{\gamma}$ the vector of loadings on the risk factor which are the regressor coefficients from the long-horizon regressions at a particular time horizon. The vector of error terms is represented by $\boldsymbol{\varepsilon}$. If the common risk factor reflected in cay explains all of the variation in and covariation between long-horizon returns, then the covariance matrix of the error terms contains only zeros.

Hence, the error term covariance matrix divided by the covariance matrix of the long-horizon returns shows how much of the actual return variation and how much of the covariation between stock markets is not explained by the common risk factor, cay. The diagonal elements of that matrix display the fraction of variation in long-horizon returns that is not explained by cay, whereas the off-diagonal elements mirror how much of the covariation between the return series is not explained by the common factor. In order to serve space I focus on the 16-quarter, local curency returns on stock indexes of the G7 economies. Table 7 presents the results. The values on the diagonal reflect that the common factor explains between 20 to 45 percent of the variation of long-horizon returns on the G7 stock indexes except Japan. As suggested by the forecast regressions almost all of the variation in returns on the Japanese stock market remain unexplained by *cay*. Hence the forecast regressions are corroborated.

The off-diagonal elements display how much of the comovement between the G7 stock markets is not explained by the common component. Not surprisingly very little of the Japanese stock market´s covariation with the other G7 economies is captured by the common risk factor. However, cay explains between one third and more than a half of the covariation between 16-quarter local currency returns on the remaining G7 stock markets. The unexplained covariation could be caused by a common permanent component in international stock markets or due to country-specific effects as suggested by Richards (1995). However, this finding highlights the importance of the common, cyclical stock market component in explaining the comovement of international stock markets at business cycle frequency.

5 Conclusions

Comovement of international stock markets on the one hand and macroeconomic explanations of stock return predictability on the other hand are individually well documented. However, a robust link between international stock markets at business cycle frequency has not been established yet.

By employing a simple manipulation of the theoretical framework proposed by Lettau and Ludvigson (2001,2004), I demonstrate that temporary fluctuations in the market value of U.S. households´ foreign equity holdings are induced by time-varying returns on foreign stock markets. Time variation of returns displays variation in risk premia. Hence, short-run fluctuations of the U.S. consumption-wealth ratio predict excess returns, a proxy for risk premia, on international stock markets. This finding suggests the existence of an important transitory component common to international stock markets that explains a considerable fraction of short-term variation in international stock returns.

Furthermore, between one third and one half of the covariation between long-horizon returns on stock markets of the G7 economies can be explained by the so far undiscovered common, temporary component which underlines its importance for the comovement of international stock markets.

Additionally, U.S. households seem to rebalance their equity portfolios in response to the perception of local currency rather than exchange rate adjusted returns as the cointegration residual explains local currency returns at least equally well as returns denominated in U.S. dollar. Moreover, exchange rate changes do not seem to cause cyclical variation in the market value of U.S. households´ foreign equity holdings which is mirrored in the non-predictablity of exchange rate changes by the cointegration residual.

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Data appendix

- U.S. household stock market wealth includes directly held equity shares at market value and indirectly held equity shares in the form of 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 such that quarterly stock market wealth has to be constructed from Flow of Funds tables L.213 and L.214.

  - Table L.213 lists 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 in stock market wealth. I calculate the amount of equities held by U.S. households through mutual fund holdings 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, I take the fraction of e.g. direct household mutual fund shares holdings at market value and multiply it with the direct holding of corporate equities by mutual funds (L.213, line 17). I apply this procedure 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 derived from 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. I assume 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 because U.S. households either directly or indirectly hold about 90% of total corporate equity issues.

- U.S. household domestic asset wealth is the 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 consumption expenditure on non-durable goods and services excluding shoes and clothing published by the Bureau of Economic Analysis in NIPA table 2.3.5. Data on total personal consumption expenditure is also taken from NIPA table 2.3.5.

- Data on U.S. labour income is freely available from the Bureau of Economic Analysis in NIPA table 2.1. I follow Lettau and Ludvigson who define labour income 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)].

- The CPI 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. is used to obtain real variables.

- The Bureau of Economic Analysis publishes population figures in NIPA table 2.1, which are used to obtain per capita figures.

- The nominal effective foreign equity investment exchange rate is a 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 sample period from first quarter 1957 to third quarter 2003. The source of bilateral U.S. dollar spot exchange rates is the IMF´s International Financial Statistic January 2004. I use the dollar-euro exchange rate for all EMU member countries under consideration since 1999.

- I define excess returns on MSCI indizes as return at the end of quarter minus the risk-free rate at the beginning of quarter, reflecting the opportunity cost of a U.S. investor investing in foreign equity. Returns are the natural logarithm of the respective index value at the end of time t+1 minus the natural logarithm of the index value at the end of time t. The risk-free rate is the 3-month-U.S. treasury bill. Since I regard logarithmic approximations of net returns (continuously compounded returns), the h-period return is the sum of the one period returns over h periods.

Tables

Table 1: Johansen Cointegration Test

	Critical Values Trace			Test Statistic Trace
	10%	5%	1%	
r=0	44.4929	47.8545	54.6815	44.1243
r=1	27.0669	29.7961	35.4628	19.0531
r=2	13.4294	15.4943	19.9349	3.6106
r=3	2.7055	3.8415	6.6349	0.0150

	Critical Values L-Max			Test Statistic L-Max
	10%	5%	1%	
r=0	25.1236	27.5858	32.7172	25.0711
r=1	18.8928	21.1314	25.8650	15.4426
r=2	12.2971	14.2639	18.5200	3.5955
r=3	2.7055	3.8415	6.6349	0.0150

	AIC	SIC
l=1	-22.0350	-21.7791
l=2	-21.9342	-21.4224

Notes: The variables under consideration are non-durables and services consumption expenditure excluding expenditures on clothing and shoes, foreign equity holdings, domestic asset wealth and labour income. All variables are measured at quarterly frequency. The sample starts second quarter 1952 and ends second quarter 2004. All variables are in natural logarithms, real, p.c. in 2000 chain weighted U.S. dollars.
The Johansen test is performed under the assumption of an unrestricted 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: VECM estimates

	Δc_t	Δfe_t	Δdaw_t	Δy_t
Δc_{t-1}	0.2412 (3.2462)	3.2414 (1.9277)	0.7594 (2.2248)	0.5450 (3.2615)
Δfe_{t-1}	−0.0032 (−1.0310)	0.0799 (1.1311)	−0.0174 (−1.2130)	0.0106 (1.5113)
Δdaw_{t-1}	−0.0048 (−0.2726)	−0.2366 (−0.5910)	0.0889 (1.0942)	−0.0601 (−1.5099)
Δy_{t-1}	0.0922 (2.5886)	0.7503 (0.9303)	0.2555 (1.5606)	−0.1192 (−1.4872)
$\hat{\boldsymbol{\beta}}'\boldsymbol{x}_{t-1}$	−0.0113 (−0.8507)	1.2974 (4.3255)	0.2254 (3.6996)	−0.0043 (−0.1445)

Notes: This table reports VECM estimates for the cointegrated VAR consisting of non-durable consumption and services consumption expenditure excluding clothing and shoes, c, foreign equity, fe, domestic asset wealth, daw, and labour income, y, for the sample period from second quarter 1952 to second quarter 2004. $\hat{\boldsymbol{\beta}}'\boldsymbol{x}_{t-1}$ is the cointegration residual. T-statistics are in parenthesis.

Table 3: Forecast error variance decompositions of the levels of the four cointegrated variables

	$\alpha_c = \alpha_y = 0$							
	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 error variance share of the level of the cointegrating variables, consumption, *c*, foreign equity, *fe*, 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 4: Forecast regressions of changes of U.S. households´ foreign equity holdings

H	$\sum_{h=1}^{H} \Delta fe_{t+h}^{\$}$	R²	$\sum_{h=1}^{H} \Delta neer_{t+h}$	R²	$\sum_{h=1}^{H} \Delta fe_{t+h}^{NC}$	R²
1	0.9701 (3.5580)	0.0509	−0.1189 (−0.8476)	-0.0028	1.0890 (3.8472)	0.0542
4	3.5296 (3.8210)	0.1986	−0.6038 (−0.9717)	0.0090	4.1334 (4.4469)	0.2226
8	6.8687 (4.7086)	0.3216	−0.5995 (−0.8416)	0.0017	7.4682 (6.0417)	0.3454
12	10.4908 (5.1706)	0.4465	−0.0606 (−0.0700)	-0.0057	10.5513 (5.4026)	0.4335
14	11.8395 (5.0127)	0.4505	0.2833 (0.2954)	-0.0048	11.5562 (5.2249)	0.4157
16	12.7522 (4.7195)	0.4292	0.5447 (0.5100)	-0.0024	12.2076 (4.8794)	0.3728
20	14.8859 (4.5463)	0.4032	1.1717 (0.8768)	0.0075	13.7143 (4.5842)	0.3108
24	16.1029 (4.3291)	0.3608	2.2608 (1.1250)	0.0437	13.8421 (3.5284)	0.2555

Notes: This table reports OLS estimates from long-horizon regressions using the cointegration residual as sole regressor. The forecast horizon *h* is in quarters. The sample spans the period from second quarter 1952 to second quarter 2004 for $\Delta fe^{\$}$ and first quarter 1957 to third quarter 2003 for $\Delta neer$ and Δfe^{NC}. R² reports values of the adjusted R². Newey-West corrected t-statistics are displayed in parenthesis.

The first column provides estimates from the regression of changes of U.S. households´ foreign equity holdings in U.S. dollar terms, $\Delta fe^{\$}$ on the cointegration residual.

The middle column gives estimates from the regression of changes of the foreign equity investment weighted nominal effective U.S. dollar exchange rate, $\Delta neer$, on the cointegration residual.

The right column presents estimates from the regression of changes of U.S. households´ foreign equity holdings in terms of a weighted basket of national currencies, Δfe^{NC} on the cointegration residual.

Table 5: Forecast regressions of excess returns on MSCI stock indexes in U.S. dollar

h	1	4	8	12	14	16	20	24
er_{AUS}	0.8712	2.8112	4.0903	5.4468	6.1099	7.1426	7.4536	7.3287
	(2.1522)	(2.3339)	(2.9489)	(2.7188)	(2.7237)	(3.1098)	(3.0703)	(3.1954)
	R^2:0.0227	R^2:0.0792	R^2:0.1073	R^2:0.1726	R^2:0.1922	R^2:0.2224	R^2:0.2244	R^2:0.2278
er_{CND}	0.4319	1.2131	2.4105	4.1117	4.8221	5.6523	6.1120	6.3098
	(1.0964)	(0.9291)	(1.4618)	(3.1971)	(3.6636)	(4.1470)	(3.9067)	(4.2549)
	R^2:0.0043	R^2:0.0120	R^2:0.0443	R^2:0.1395	R^2:0.1967	R^2:0.2530	R^2:0.2181	R^2:0.2551
er_{FIN}	0.6363	2.8908	10.5780	20.2019	24.1283	28.5109	30.6139	32.2378
	(0.4569)	(0.6130)	(2.0124)	(8.2004)	(8.2669)	(6.4269)	(3.7237)	(3.8186)
	R^2:-0.0056	R^2:0.0121	R^2:0.1199	R^2:0.3059	R^2:0.3760	R^2:0.4152	R^2:0.3529	R^2:0.2161
er_{FRA}	0.9975	3.4430	5.9605	9.8685	11.4769	12.7556	12.7983	12.6190
	(2.1719)	(2.2504)	(2.9405)	(4.4095)	(4.9087)	(5.1132)	(4.8022)	(3.9669)
	R^2:0.0303	R^2:0. 0919	R^2:0.1421	R^2:0.2982	R^2:0. 3523	R^2:0.3699	R^2:0.3345	R^2:0.2161
er_{GER}	1.0284	2.8656	4.9030	6.8826	7.2341	7.4094	5.6867	4.7148
	(2.6066)	(2.3438)	(2.0695)	(2.1946)	(2.1172)	(1.8927)	(1.4870)	(1.3976)
	R^2:0.0397	R^2:0.0810	R^2:0.1323	R^2:0.1955	R^2:0.1936	R^2:0.1735	R^2:0.0831	R^2:0.0551
er_{HK}	0.6945	2.9609	5.9392	7.0278	7.5809	7.3780	5.8164	4.1227
	(0.8630)	(1.0041)	(1.7642)	(2.1702)	(2.5882)	(2.4916)	(1.9486)	(1.1030)
	R^2:-0.0001	R^2:0.0282	R^2:0.0692	R^2:0.0897	R^2:0.1151	R^2:0.1021	R^2:0.0546	R^2:0.0188
er_{IRL}	1.0663	3.9169	7.7882	9.9650	11.0006	13.7069	15.4923	12.9429
	(2.6858)	(4.5234)	(4.9633)	(3.9062)	(3.6189)	(4.7835)	(7.9547)	(6.4231)
	R^2:0.0458	R^2:0.2107	R^2:0.4692	R^2:0.4517	R^2:0.4431	R^2:0.4209	R^2:0.3918	R^2:0.1275
er_{ITA}	1.0826	4.2690	8.5645	13.4730	15.6659	17.7489	19.7154	21.0160
	(2.1476)	(2.3288)	(3.0344)	(4.1737)	(4.6330)	(5.4904)	(6.8876)	(7.5571)
	R^2:0.0311	R^2:0.1102	R^2:0.2066	R^2:0.3611	R^2:0.4267	R^2:0.4744	R^2:0.5047	R^2:0.5162
er_{JPN}	0.3045	0.8852	2.7827	4.5474	4.9615	4.3721	0.6934	−2.9349
	(0.5280)	(03762)	(0.7124)	(0.9037)	(0.9020)	(0.7402)	(0.1153)	(−0.4645)
	R^2:-0.0038	R^2:-0.0019	R^2:0.0167	R^2:0.0391	R^2:0.0421	R^2:0.0262	R^2:-0.0079	R^2:0.0022
er_{KOR}	−0.3517	−4.2011	−7.3567	−10.6990	−12.0323	−12.0158	−19.6171	−33.3904
	(−0.2309)	(−08925)	(−1.5213)	(−2.0172)	(−2.2734)	(−2.3204)	(−4.1076)	(−2.8768)
	R^2:-0.0147	R^2:0.0233	R^2:0.0655	R^2:0.1534	R^2:0.1796	R^2:0.1539	R^2:0.3274	R^2:0.4642
er_{MEX}	0.6169	1.4343	2.5941	2.7141	0.6505	−1.6748	−7.4272	−20.3085
	(0.6529)	(0.5661)	(0.6003)	(0.5658)	(0.1261)	(−0.2460)	(−0.8722)	(−1.3167)
	R^2:-0.0093	R^2:-0.0081	R^2:-0.0027	R^2:-0.0092	R^2:-0.0195	R^2:-0.0184	R^2:0.0089	R^2:0.1164

Table 5 continued

h	1	4	8	12	14	16	20	24
er_{NL}	1.0005	3.5965	6.7214	9.1531	9.9173	10.5030	9.5943	8.5472
	(3.1157)	(3.6089)	(4.6815)	(4.4616)	(4.1096)	(3.6327)	(3.3981)	(3.1030)
	R^2:0.0539	R^2:0.2045	R^2:0.3505	R^2:0.4449	R^2:0.4380	R^2:0.4084	R^2:0.2849	R^2:0.2203
er_{SIN}	0.0440	− 0.2948	− 0.1769	0.4696	0.8497	0.9206	−1.9408	− 4.5948
	(0.0579)	(−0.0987)	(−0.0434)	(0.1142)	(0.2270)	(0.2681)	(−0.5987)	(−1.2408)
	R^2:-0.0074	R^2:-0.0071	R^2:-0.0077	R^2:-0.0075	R^2:-0.0063	R^2:-0.0062	R^2:-0.0003	R^2:0.0354
er_{ESP}	0.4877	2.2619	5.8918	10.1523	12.7383	16.0898	21.0779	26.4739
	(1.1275)	(1.2738)	(1.6150)	(2.0498)	(2.3582)	(2.8577)	(3.9615)	(5.4115)
	R^2:0.0015	R^2:0.0305	R^2:0.0879	R^2:0.1513	R^2:0.1979	R^2:0.2620	R^2:0.3514	R^2:0.4582
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
er_{US}	0.7971	3.0509	6.1795	8.7082	9.6041	10.8230	11.3735	12.3250
	(2.8128)	(3.2770)	(5.1427)	(5.3560)	(5.2413)	(5.0788)	(5.1519)	(6.4627)
	R^2:0.0446	R^2:0.1798	R^2:0.3802	R^2:0.5060	R^2:0.5291	R^2:0.5821	R^2:0.5202	R^2:0.5351

Notes: This table reports OLS estimates from forecast regressions of excess returns on Morgan Stanley Capital International (MSCI) stock indexes with underlying market capitalisation in current U.S. dollar on the cointegration residual as 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 the end of 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, reflecting the opportunity cost of foreign stock market investment for a U.S. investor and r_t the end-of period return. As logarithmic approximations are employed, the h-period excess return is the sum of one period excess returns over h periods.

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, United Kingdom and the United States.

Table 6: Forecast regressions of excess returns on MSCI stock indexes in local currency

h	1	4	8	12	14	16	20	24
er_{AUS}	0.9266	3.2216	5.2512	7.0189	7.5932	8.3696	8.4029	7.7399
	(2.6657)	(2.9836)	(3.4087)	(3.3911)	(3.5216)	(3.6840)	(3.8692)	(3.4903)
	R^2:0.0384	R^2:0.1343	R^2:0.2090	R^2:0.2968	R^2:0.3033	R^2:0.3174	R^2:0.2942	R^2:0.2627
er_{CND}	0.4660	1.4304	2.9540	4.9892	5.6243	6.2323	6.0980	5.4482
	(1.2743)	(1.1919)	(2.0095)	(4.3010)	(4.7550)	(4.8273)	(3.2416)	(2.3623)
	R^2:0.0091	R^2:0.0247	R^2:0.0783	R^2:0.2000	R^2:0.2434	R^2:0.2804	R^2:0.2162	R^2:0.1826
er_{FIN}	0.3107	2.0966	10.3794	20.8109	24.5590	28.3636	29.8284	30.1387
	(0.2250)	(0.4324)	(1.7702)	(8.1545)	(10.0164)	(7.6492)	(4.0928)	(2.9690)
	R^2:-0.0102	R^2:0.0003	R^2:0.1103	R^2:0.3072	R^2:0.3647	R^2:0.3872	R^2:0.3167	R^2:0.1700
er_{FRA}	0.8956	3.2478	6.2456	10.5215	10.7018	12.1353	13.4310	13.2312
	(2.3809)	(2.4397)	(3.6279)	(6.2620)	(6.7411)	(7.5508)	(6.8853)	(7.2310)
	R^2:0.0280	R^2:0.1043	R^2:0.2073	R^2:0.4440	R^2:04970	R^2:0.5127	R^2:0.4905	R^2:0.4471
er_{GER}	0.9928	2.8814	5.4218	8.1692	9.4555	9.9331	8.7733	9.0559
	(3.0440)	(2.8861)	(2.5626)	(2.7923)	(4.0941)	(4.1620)	(2.5125)	(3.3135)
	R^2:0.0380	R^2:0.0891	R^2:0.1734	R^2:0.2900	R^2:0.3552	R^2:0.2915	R^2:0.2319	R^2:0.2440
er_{HK}	0.8295	3.4348	6.7870	8.2873	8.9803	8.9488	7.3807	5.3590
	(1.0455)	(1.1671)	(2.0253)	(2.5952)	(3.0535)	(2.9852)	(2.2611)	(1.2707)
	R^2:0.0038	R^2:0.0429	R^2:0.1002	R^2:0.1446	R^2:0.1856	R^2:0.1724	R^2:0.1053	R^2:0.0439
er_{IRL}	0.6833	2.8015	7.8372	11.8549	13.1840	16.1804	19.3477	20.2432
	(1.7075)	(2.7008)	(4.4004)	(6.0859)	(5.4869)	(6.6709)	(11.1509)	(7.2006)
	R^2:0.0036	R^2:0.0772	R^2:0.3672	R^2:0.5256	R^2:0.5384	R^2:0.5910	R^2:0.5238	R^2:0.1275
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

Table 6 continued

h	1	4	8	12	14	16	20	24
er_{NL}	0.9602 (3.3022) R^2:0.0505	3.5813 (3.6431) R^2:0.1847	7.1335 (4.8319) R^2:0.3484	10.2756 (4.6533) R^2:0.4864	11.1911 (4.4992) R^2:0.4884	11.9504 (4.1560) R^2:0.4813	12.2255 (4.0722) R^2:0.4218	12.1612 (4.5859) R^2:0.3819
er_{SIN}	0.0736 (0.1025) R^2:-0.0073	– 0.1427 (–0.0508) R^2:-0.0074	0.2467 (0.0680) R^2:-0.0077	1.0871 (0.3220) R^2:-0.0039	1.5117 (0.5178) R^2:0.0003	1.7224 (0.6636) R^2:0.0024	– 0.4926 (–0.2022) R^2:-0.0078	– 2.6078 (–0.9897) R^2:0.0143
er_{ESP}	0.5208 (1.4210) R^2:0.0036	2.4722 (1.7083) R^2:0.0501	6.7432 (2.3969) R^2:0.1676	11.5497 (3.1202) R^2:0.2828	13.7213 (3.4419) R^2:0.3226	16.4466 (3.9288) R^2:0.3785	20.6856 (5.4473) R^2:0.4514	25.0703 (7.7154) R^2:0.5383
er_{SWE}	0.9363 (2.1320) R^2:0.0235	2.7052 (1.9210) R^2:0.0482	6.0884 (2.5817) R^2:0.1476	10.4780 (3.7663) R^2:0.3115	11.9249 (4.2812) R^2:0.3460	13.0336 (4.7144) R^2:0.3650	13.7147 (7.4838) R^2:0.3470	14.6489 (7.9685) R^2:0.3457
er_{CH}	0.8786 (2.9600) R^2:0.0408	3.3948 (3.4649) R^2:0.1759	6.5848 (4.7978) R^2:0.3220	9.6324 (5.4072) R^2:0.4723	10.4105 (4.9490) R^2:0.4665	11.5120 (4.6089) R^2:0.4754	12.3286 (4.0036) R^2:0.4069	13.1869 (4.1445) R^2:0.3984
er_{UK}	1.0514 (3.0164) R^2:0.0544	3.6953 (2.6453) R^2:0.1703	6.6144 (4.1872) R^2:0.3271	8.8991 (5.2282) R^2:0.4269	8.9403 (4.7869) R^2:0.4075	9.2532 (4.1870) R^2:0.4057	7.5599 (4.0982) R^2:0.2866	5.8112 (4.3172) R^2:0.1789
er_{US}	0.7971 (2.8128) R^2:0.0446	3.0509 (3.2770) R^2:0.1798	6.1795 (5.1427) R^2:0.3802	8.7082 (5.3560) R^2:0.5060	9.6041 (5.2413) R^2:0.5291	10.8230 (5.0788) R^2:0.5821	11.3735 (5.1519) R^2:0.5202	12.3250 (6.4627) R^2:0.5351

Notes: This table reports OLS estimates from forecast regressions of excess returns on Morgan Stanley Capital International (MSCI) stock indexes with underlying market capitalisation in local currency on the cointegration residual as 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 the end of 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, , reflecting the opportunity cost of foreign stock market investment for a U.S. investor and r_t the end-of-period return. As logarithmic approximations are employed, the h-period excess return is the sum of one period excess returns over h periods.

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, United Kingdom and the United States.

Table 7:Covariance matrix

	CND	FRA	GER	ITA	JPN	UK	US
CND	0.8108						
FRA	0.6660	0.5993					
GER	0.6550	0.6185	0.7258				
ITA	0.6625	0.5310	0.5278	0.5965			
JPN	0.8689	0.8319	0.8225	0.8466	0.9773		
UK	0.7139	0.6283	0.6718	0.5716	0.8656	0.7461	
US	0.6380	0.5280	0.6079	0.4468	0.7410	0.6121	0.5615

Notes: This table provides the covariance matrix of the error terms divided by the covariance matrix of actual 16-quarter returns on the G7 MSCI stock indexes from

$$\mathrm{var}(r) = \gamma\gamma' \, \mathrm{var}(cay) + \mathrm{cov}(\varepsilon)$$

$$1 = \frac{\gamma\gamma' \, \mathrm{var}(cay)}{\mathrm{var}(r)} + \frac{\mathrm{cov}(\varepsilon)}{\mathrm{var}(r)}$$

where $\boldsymbol{r}$ is the vector of 16-quarter returns on the G7 indexes, ***cay*** represents the cointegration residual, i.e. the common risk factor and $\boldsymbol{\gamma}$ the vector of loadings on the risk factor which are the regressor coefficients from the long-horizon regressions at 16 quarter horizon and $\boldsymbol{\varepsilon}$ denotes the vector of error terms.
Elements on the diagonal display how much of the variance of the 16-quarter G7 excess returns denominated in national currency remains unexplained. The off-diagonal elements measure how much of the covariation between the G7 returns is not explained by ***cay***.

Figures

Figure 1 (a)

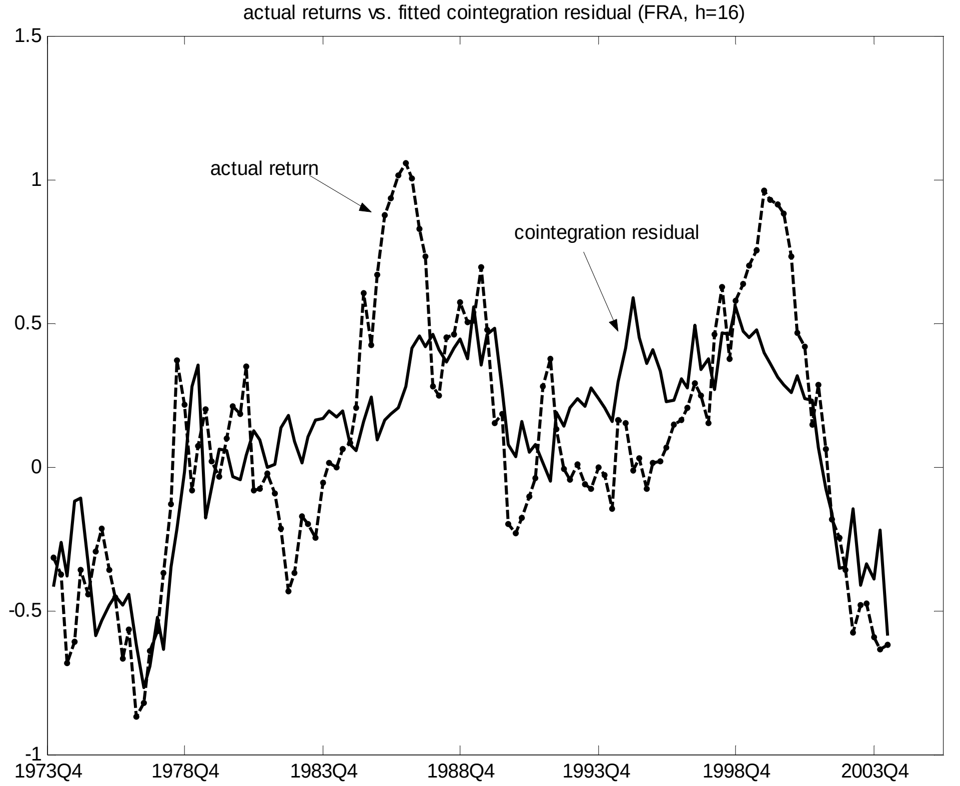


Figure 1 (b)

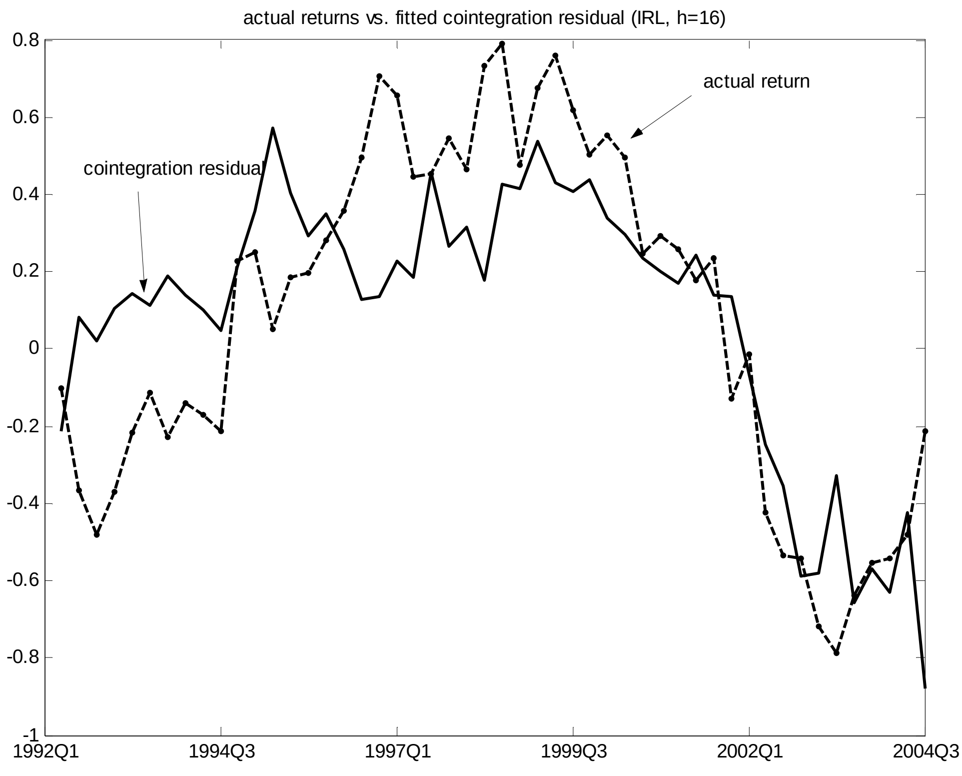


Figure 1 (c)

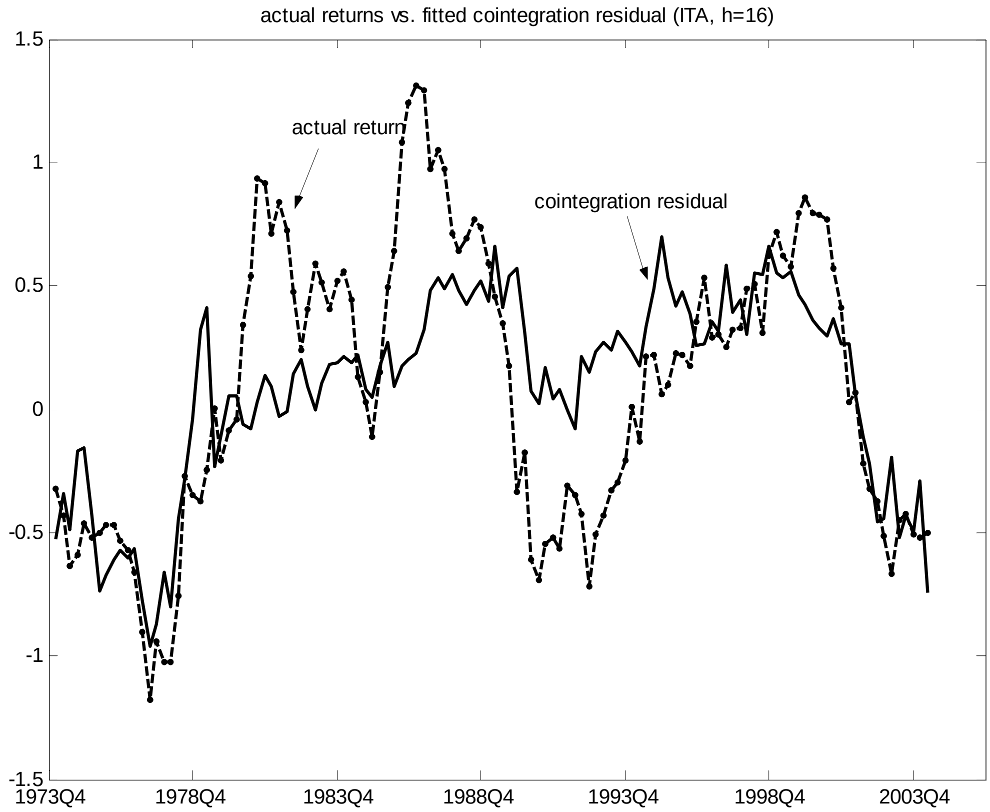


Figure 1 (d)

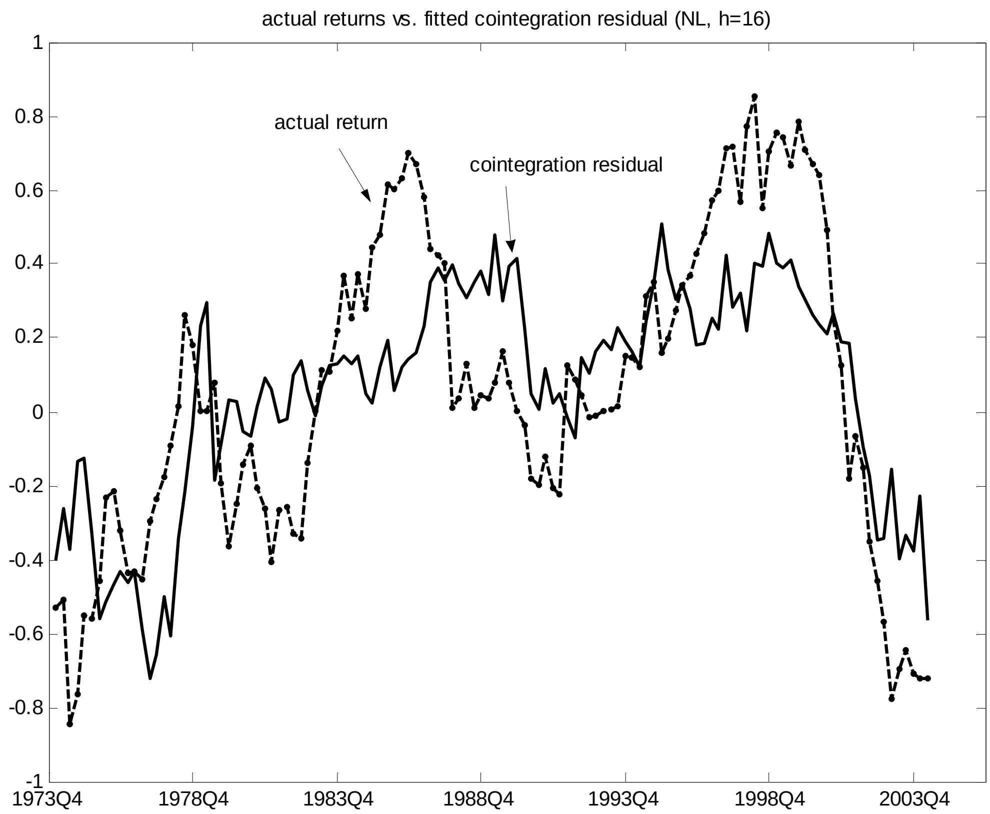


Figure 1 (e)

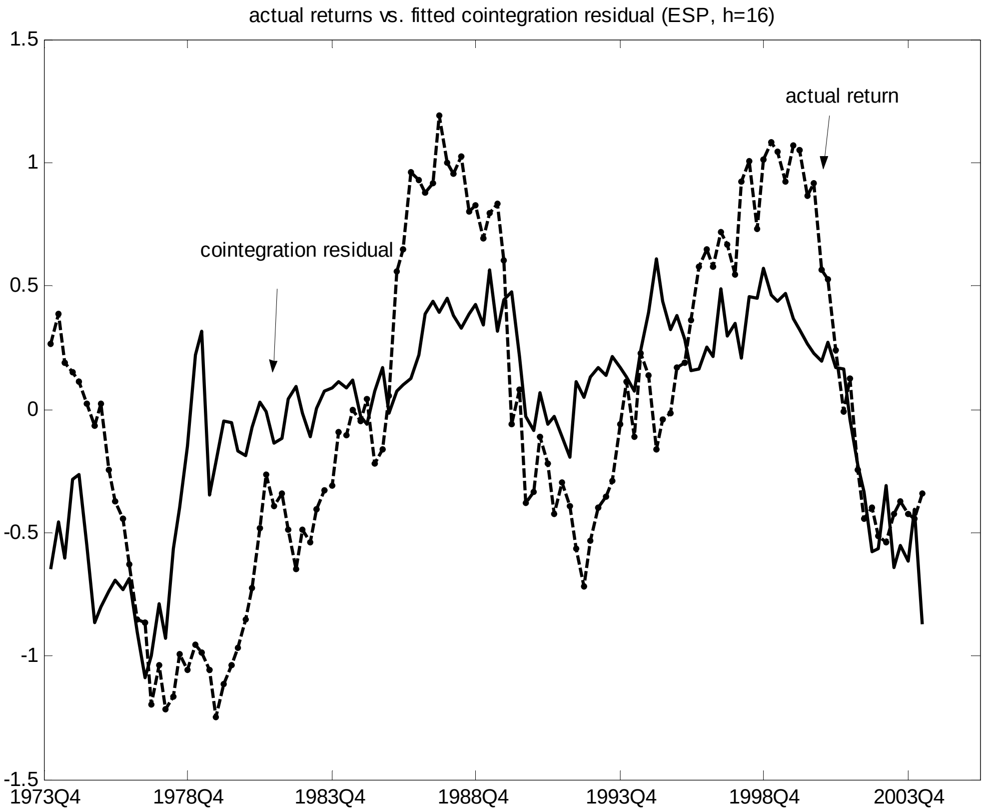


Figure 1 (f)

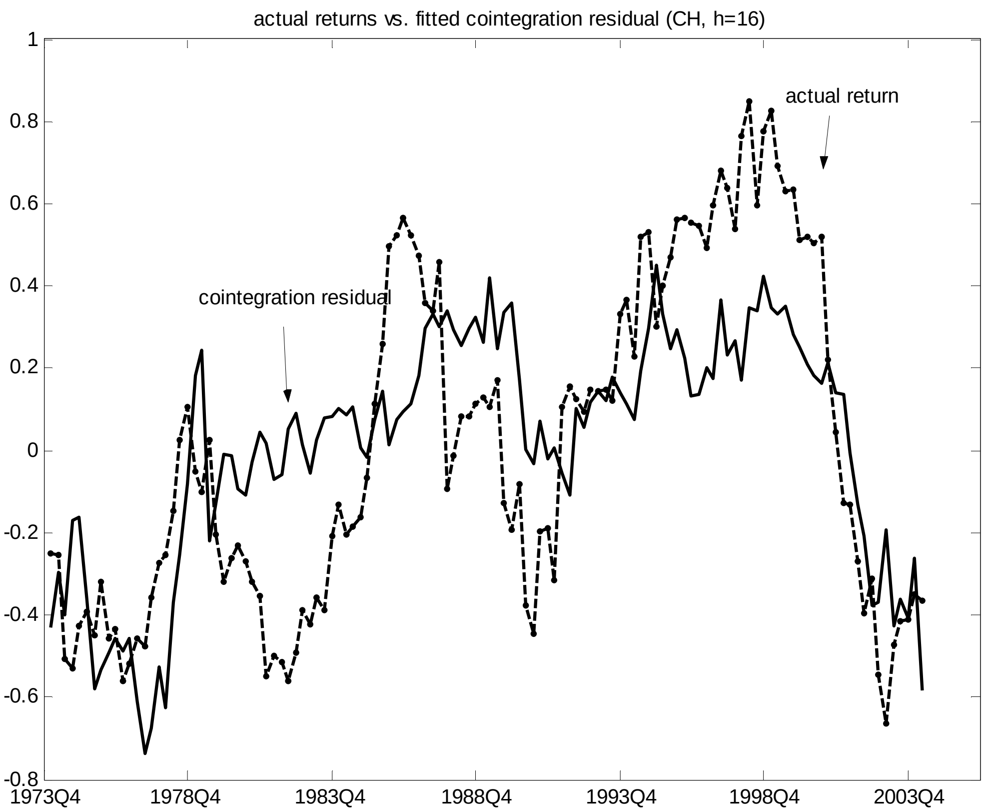


Figure 1 (g)

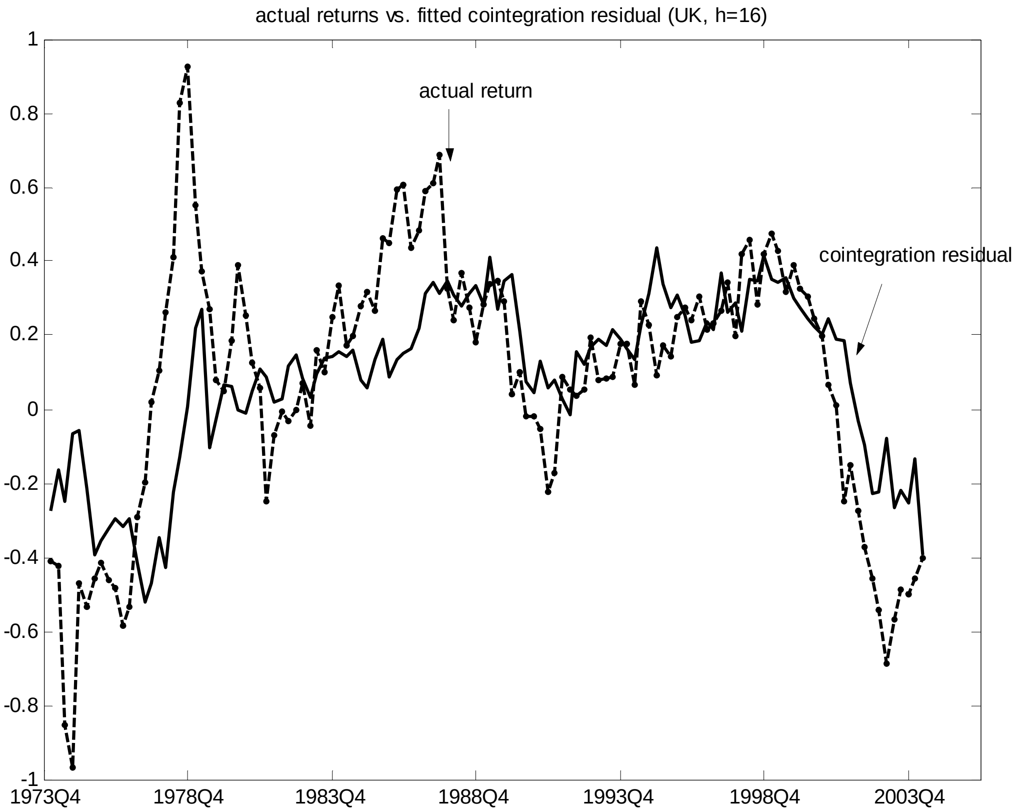
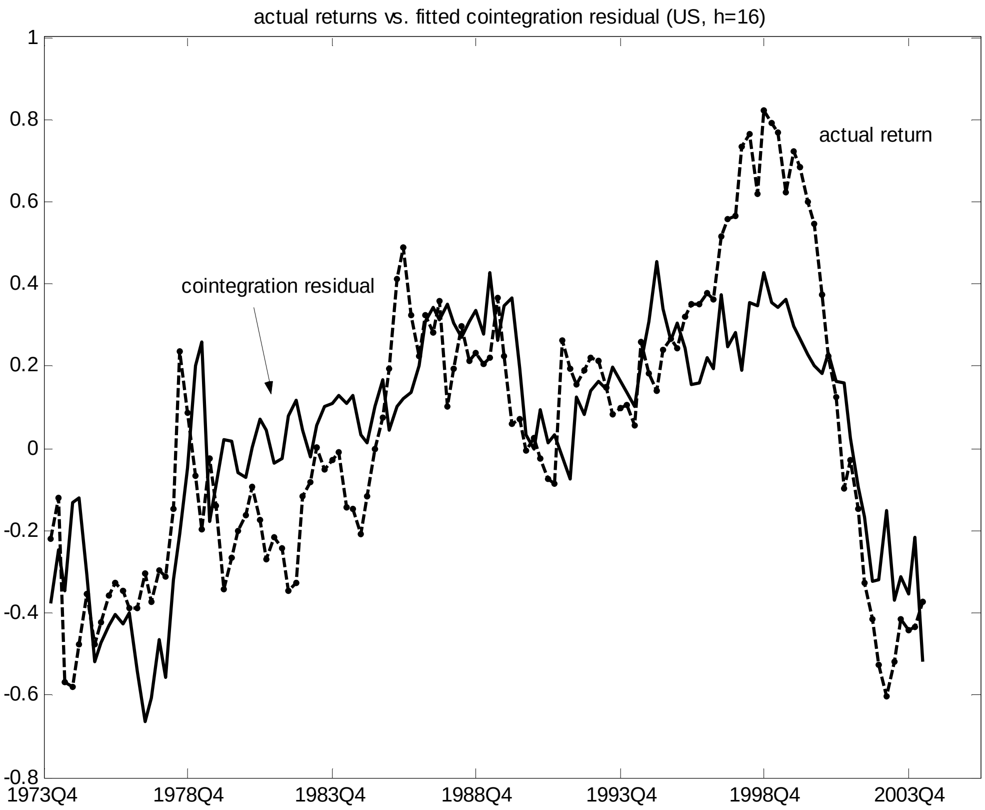


Figure 1 (h)



Figures 1 (a) to (h) caption:

The figures 1(a) to (h) display realisations of 16-quarter returns of MSCI indexes (dashed line) together with the fitted value of the cointegration residual (straight line) at that time horizon. The country indexes in question are France, Ireland, Italy, Netherlands, Spain, Switzerland, United Kingdom as well as the United States. The sample starts 1973 for all countries except Ireland