Global asset pricing: Is there a role for long-run consumption risk?

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February 2009

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Abstract

We estimate long-run consumption-based asset pricing models using a comprehensive set of international test assets: country indices, international value/growth portfolios, international bond portfolios, and FX portfolios. We find that long-run risk considerably reduces the estimates of the risk-aversion parameter, compared to cases where we use contemporaneous consumption growth to estimate the risk aversion parameter. We also find, however, that long-run risk models only price the cross-sectional/cross-country variation in returns better when we allow for a common constant mispricing of the assets, i.e. include a constant in the empirical moment function. We investigate why certain results differ from those reported in the literature that uses U.S. data.

Keywords: International Asset Pricing, Long-run consumption risk

JEL-classification: F30, G12, G15

1 Introduction

The main insight of the consumption-based asset pricing model is that differences in returns across assets can be explained by the assets' exposure to consumption risk. In its standard form, this model has failed on a grand scale (see, e.g., Breeden, Gibbons & Litzenberger, 1989 or Mankiw & Shapiro, 1996). In a more recent version, though, where consumption growth is allowed to contain a small predictable component, such that returns are determined by their exposures to long-run consumption growth, the model's performance improves considerably when tested on cross-sections of U.S. assets; see Daniel & Marshall (1997), Parker (2001, 2003), Bansal & Yaron (2004), Parker & Julliard (2003, 2005), Hansen, Heaton & Li (2008), and Malloy, Moskowitz & Vissing-Jørgensen (2008).

In this paper, we use a very broad set of international test assets (broad stock indices, value/growth stock indices, bonds, and currencies) to evaluate whether asset-pricing models in which long-run consumption risk to plays a role (LRR-CAPMs) capture cross-country differences in returns better than models that use contemporaneous consumption growth as a risk factor. Our main result is surprisingly consistent across test assets: The risk-aversion coefficient of the representative consumer is estimated at a more reasonable level when we allow for long-run risk. We also find that long-run risk enables the models to better capture the cross-country variation in returns – however, only when we allow for a common mispricing of all assets (a constant in the empirical moment function).

In our empirical specification, we follow Malloy, Moskowiz & Vissing-Jørgensen (2008), who themselves are inspired by Hansen, Heaton & Li (2008). The models in Malloy et al. (2008) and Hansen et al. (2008) are based upon the recursive utility framework. An attractive feature of this framework is that it allows for a separation of the coefficient of relative risk aversion from the elasticity of intertemporal substitution in consumption (EIS). Like in Malloy et al. (2008), we focus on the situation where the EIS is equal to one. When the EIS is equal to one, the stochastic discount factor (SDF) turns out to be a simple function of discounted future growth rates of consumption. Indeed, when the EIS is equal to one, the SDF is very similar to the SDF promoted by Parker & Julliard (2005).

We test the model on excess returns from four sets of international test assets: Returns from country equity indices (we use returns from the G-7 equity markets), returns from international value/growth portfolios, returns from international bond portfolios, and returns from international FX carry trade portfolios. The sample period coincides with the post-Bretton Woods period. Our use of many different sets of test assets help us getting around the Lewellen, Nagel & Shanken (2008) insight that factor-pricing models sometimes work on few specific data sets only.

When testing a consumption-based asset pricing model using U.S. returns, one uses U.S. consumption. But what measure of consumption should be used when the test assets are international? As our test assets are international G-7 assets, we use the series for G-7 consumption growth, calculated by the IMF, as our measure of the consumption-risk factor. When using G-7 consumption, we implicitly assume the existence of internationally integrated capital markets, as internationally integrated capital markets help smoothing out idiosyncratic fluctuations in consumption, such that only the risk of fluctuations in aggregate consumption is rewarded in equilibrium. Using a newly developed measure of segmentation, Bekaert, Harvey, Lundblad & Siegel (2008) show that the segmentation of developed equity markets is not larger than the segmentation of the U.S. equity market has been since the late 1980s, i.e. it seems reasonable to use an international consumption series. In addition, to make sure that our results are not driven by our choice of consumption series, we conduct two robustness tests: We estimate our models on a subsample in which the developed equity markets are as integrated as the U.S. capital market is, i.e. a subsample starting in the late 1980s, and we estimate the models using U.S. consumption instead of G-7 consumption. Qualitatively, we find the same results in these robustness tests as in our baseline case.[1]

Overall, we perform two kinds of estimations: One where we allow for a constant in the empirical moment function and one where we do not. We find that the estimate of the risk aversion coefficient declines with the horizon over which the growth in consumption is measured when we do not include a constant in the empirical moment function. For instance, when we estimate the model using the returns from the G-7 equity markets as test assets, we find an estimate of the risk aversion coefficient of 117.84 when consumption

[1] Early papers using U.S. consumption to test international consumption-based asset pricing models include Cumby (1990) and Wheatley (1988), whereas papers that use world-consumption risk to price international assets include Sarkissian (2003) and Li & Zhong (2004).

growth is measured over one quarter (the standard consumption CAPM for an investor with power utility). However, when the stochastic discount factor includes growth in consumption measured over the next 12 quarters, the estimate of the risk aversion decreases by approximately 75%, to reach a value of 28.97. In other words, when we do not include a constant in the empirical moment function, the estimate of the risk aversion coefficient is drastically reduced when allowing for long-run risk.

We also estimate models where we allow for a constant in the SDF. The reason why it has become common practice (see, e.g., Parker & Julliard, 2005 and MMVJ, 2008) to allow for a constant, even if it should theoretically not be there, is that it allows for an evaluation of the extent to which the model under- or overpredicts the return on the different test assets, i.e. the extent to which there is an equity premium. It turns out that the choice of including a constant or not is important in our sample with international test assets: When there is a constant in the moment function, long-run risk does not have an effect on the estimate of the risk-aversion parameter.

What about the extent to which consumption growth risk can account for the cross-country differences in returns? If no constant is included in the empirical moment function, the cross-sectional R^2 is often estimated to be negative, regardless of the horizon over which consumption growth is measured. On the other hand, the R^2s increase with the horizon when the constant is included. For instance, the cross-sectional R^2 is virtually zero when using one-period growth in consumption (the standard model) to price the return from the broad market indices, but increases to 26% when we include growth in consumption measured over the next 12 quarters.

Hence, there is a certain tension in our results: If a constant is not included in the empirical moment function, the risk-aversion coefficient is estimated at more reasonable levels when allowing for long-run risk, but the cross-sectional fit is not improved; if there is constant, allowing for long-run risk improves the cross-sectional fit, but the estimate of the risk-aversion coefficient is imprecise and sometimes even negative. These findings are common to the broad range of test assets we use.[2] In contrast to results based upon

[2] Discussions of the differences that arise between models that include a constant and models that do not have recently taken center stage in investigations of currency risk premia. Burnside (2007) argues that it is of great importance whether a constant is included in the moment function when evaluating whether

U.S. data, though, long-run risk does not help in both dimensions at the same time when testing the model on international data. We show one reason why this could be so: Long-run G-7 consumption growth is not as predictable by international asset returns as U.S. consumption growth is reported to be in Parker & Julliard (2005) and MMVJ (2008).

There are by now many papers that use long-run risk models to price U.S. assets. However, to the best of our knowledge, only few papers exist that use long-run risk models to investigate different characteristics of international financial markets. For instance, Bansal & Shaliastovich (2008) use a long-run risk model to explain deviations from the expectation hypothesis in bond and currency markets. The difference to our paper is that we focus on capturing the cross-country variation in returns from different international assets markets. Colacito & Croce (2006) investigate the gains from international portfolios diversification in a long-run risk model. Colacito & Croce calibrate their model and show that the welfare gains from international financial integration are larger than those normally reported in the literature, primarily because investors in their model face long-run consumption risk and are able to share such long-run risk in a financially integrated world. Colacito & Croce do not estimate their model, though, and all-in-all focus on other issues than we do. Rangvid (2008) shows, like in this paper, that the estimate of the risk-aversion coefficient is lower when the horizon over which consumption growth is measured is longer, when there is no constant in the empirical moment function. Rangvid does not investigate what happens if a constant is included in the moment function, though. In addition, Rangvid only tests on broad international equity indices, i.e. he does not study value/growth portfolios, bond portfolios, or FX portfolios.

The remaining part of the paper is structured as follows: In the next section, we describe the theoretical model that which we consequently take to the data. In section 3, the data are presented. Section 4 contains the main empirical results of the paper. In section 5, we explain what drives our results by evaluating the extent to which the returns we use predict future consumption growth. In section 6 we present robustness checks on our basic findings. Section 7 concludes.

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a consumption based asset pricing model is able to explain the return to currency risk premia. Lustig & Verdelhan (2008), on the contrary, argue that the constant is not important.

2 Theoretical framework and empirical implementation

Our estimations are based on the empirical framework of Malloy, Moskowiz & Vissing-Jørgensen (2008), henceforth MMVJ, that in turn builds on the theoretical work of Hansen, Heaton & Li (2008), henceforth HHL. HHL and MMVJ assume that investors have recursive Epstein & Zin (1989) preferences:

$$V_t = \left[(1-\beta) C_t^{1-\frac{1}{\rho}} + \beta \left[E_t \left(V_{t+1}^{1-\gamma} \right) \right]^{\frac{1-\frac{1}{\rho}}{1-\gamma}} \right]^{\frac{1}{1-\frac{1}{\rho}}} \tag{1}$$

where C_t is the level of the investor's consumption, ρ the elasticity of intertemporal substitution in consumption, γ the coefficient of relative risk-aversion, and β the rate of time preferences. HHL and MMVJ also assume that log-consumption dynamics are not *iid*, but move over time according to a first-order VAR driven by a unspecified vector of stationary state variables:

$$\begin{aligned} \Delta c_{t+1} &= \mu^c + U_c x_t + \lambda_0 \omega_{t+1} \\ x_{t+1} &= G x_t + H \omega_{t+1} \end{aligned} \tag{2}$$

where $\Delta c_{t+1} = \ln(C_{t+1}) - \ln(C_t)$, x_t a vector of state variables, and G has eigenvalues less than one. Under these assumptions, HHL and MMVJ show that the conditional log-linearized asset-pricing (Euler) equation for the return on an asset i over and above the risk-free rate is:[3]

$$E_t(r_{t+1}^i) - r_{t+1}^f + 0.5\sigma_t^2(r_{t+1}^i) \simeq (\gamma - 1)\, cov_t \left(r_{t+1}^i, \sum_{s=0}^{\infty} \beta^s \Delta c_{t+1+s} \right) \tag{3}$$

when the elasticity of intertemporal substitution in consumption is equal to one. In Eq. (3), $r_{t+1}^i = \ln(1+R_{t+1}^i)$ is the log-return on asset i, $r_{t+1}^f = \ln(1+R_{t+1}^f)$ is the log risk-free rate, and $0.5\sigma_t^2(r_{t+1}^i)$ is the usual variance term arising from the log-linearization.

The key point to notice in Eq. (3) is that *one-period* excess returns are determined by a *sum of discounted future consumption growth rates.* It is particularly this feature of the asset-pricing equation that sets Eq. (3) apart from the standard (power utility) asset-pricing equation relating one-period returns to one-period growth rates of consumption.[4]

[3] Appendix A reviews the main steps leading to Eq. (3).

[4] With power utility, excess returns are determined by $E_t \left[(C_{t+1}/C_t)^{-\gamma} (R_{t+1}^i - R_{t+1}^f) \right] = 0$ and the conditional log-linearized version is $E_t(r_{t+1}^i) - r_{t+1}^f + 0.5\sigma_t^2(r_{t+1}^i) = \gamma cov_t \left[\ln(C_{t+1}/C_t), \ln(R_{t+1}^i) \right]$.

In other words, the framework of HHL and MMVJ allows long-run consumption risk to play a role when determining asset prices.

Eq. (3) is based on the assumption of an elasticity of intertemporal substitution (EIS) equal to one. Given that $\rho = 1$ but γ will be estimated, the framework allows for a separation of the EIS from the risk-aversion coefficient, which the standard power utility framework does not; with power utility $\gamma = 1/\rho$, i.e. in the standard framework, the EIS is forced to equal the reciprocal of the risk aversion coefficient. MMVJ discuss why it is reasonable to work under the assumption of $\rho = 1$ when estimating the risk-aversion coefficient from a cross-section of assets, as we do in this paper.

MMVJ estimate Eq. (3) using the 25 Fama & French U.S. equity portfolios as tests assets. MMVJ find that γ is estimated to be high and/or γ is imprecisely estimated when only a few future consumption growth rates are included in the asset pricing equation. However, when allowing long-run risk to play a role, i.e. when including several future consumption growth rates in the empirical asset pricing equation, γ is estimated to be low and/or more precisely estimated.[5] In addition, they find that the extent to which the cross-sectional variation in returns can be captured by the model is often higher when allowing for long-run consumption risk.

2.1 Comparison with Parker & Julliard

The asset-pricing equation of Eq. (3) is approximate, as it is a log-linearized one. Parker & Julliard (2005) estimate an exact (but non-linear) asset-pricing relation that also relates one-period excess returns to multiperiod growth rates of consumption. Hence, it is instructive to compare Eq. (3) with the asset-pricing equation from Parker & Julliard (2005), henceforth PJ.

PJ note that the Euler-equation for the risk-free rate between any two time points $t+1$ and $t+S$, with S possibly larger than 1, is given by $U'(C_{t+1}) = \beta E_{t+1}\left[U'(C_{t+S})R^f_{t+1,t+S}\right]$. PJ substitute this expression for $U'(C_{t+1})$ into the general Euler equation for the excess

[5] MMVJ also focus on the differences in results that are obtained when using consumption of stockholders and non-stockholders, respectively. MMVJ find that the consumption of stockholders reacts more to asset returns, and consequently that γ is estimated to be lower for stockholders. MMVJ also look at other U.S. assets (e.g. bond returns).

return on any asset i between periods t and $t+1$, which results in

$$E_t \left[m_{t+S}^S \left(R_{t+1}^i - R_{t,t+1}^f \right) \right] = 0, \tag{4}$$

where $m_{t+S}^S = (C_{t+S}/C_t)^{-\gamma} R_{t+1,t+S}^f$ after assuming that the functional form for $U(C_t)$ is the standard power-utility function $U(C_t) = C_t^{1-\gamma}/(1-\gamma)$.

Eq. (4) relates multiperiod (S-period) consumption growth to one-period returns on equity followed by a $S-1$ period return from the risk-free asset. If the variation in the risk-free rate is not too big, this essentially means that one-period excess returns are related to multiperiod consumption growth rates, as in MMVJ. PJ test their model on the 25 Fama & French portfolios. Like MMVJ, Parker & Julliard find that when several future consumption growth rates are included in the empirical asset pricing equation (i.e. when $S > 1$), γ is estimated to be lower and/or more precisely estimated, compared to the standard situation where $S = 1$. Parker (2001, 2003) uses similar approaches to show that γ is estimated to be lower when S is larger.

Given the results of Parker, Parker & Julliard, and MMVJ, there is thus substantial empirical evidence that long-run risk is important when pricing U.S. assets. This clearly constitutes a valuable insight. Likewise, one might also expect long-run risk to be an important determinant for pricing international assets. In this paper we test whether this is the case.

2.2 Empirical implementation

We estimate Eq. (3) as it allows for a separation of the risk-aversion coefficient from the EIS, which Eq. (4) does not. As a robustness check we also estimate Eq. (4), though.

When empirically implementing the model, we take unconditional expectations on both sides of Eq. (3). The empirical moment function with $f_{t+1} = \sum_{s=0}^{S-1} \beta^s \Delta c_{t+1+s}$, i.e. after having truncated the infinite sum of future consumption growth rates in Eq. (3) at horizon S, then reads:

$$h\left(\gamma, \alpha, \mu_{c,S}; \Theta_{t+1}\right) = \begin{bmatrix} \mathbf{r}_{t+1}^e + 0.5\boldsymbol{\sigma}^2 - \alpha\boldsymbol{\iota}_N - (\gamma - 1)\,\mathbf{r}_{t+1}^e \left(f_{t+1} - \mu_{c,S}\right) \\ f_{t+1} - \mu_{c,S} \end{bmatrix} \tag{5}$$

where $\mathbf{r}_{t+1}^e$ is the vector of excess returns from the N test assets and $\boldsymbol{\sigma}^2$ collects the variances:

$$\mathbf{r}^e_{t+1} = \begin{bmatrix} r^1_{t+1} - r^f_{t+1} \\ \vdots \\ r^N_{t+1} - r^f_{t+1} \end{bmatrix} \quad \text{and} \quad \boldsymbol{\sigma}^2 = \begin{bmatrix} var(r^1_{t+1}) \\ \vdots \\ var(r^N_{t+1}) \end{bmatrix}.$$

In our empirical implementation, we assume conditionally homoskedastic returns, i.e. we use the unconditional return variance (computed over the full sample) as our estimate of $\boldsymbol{\sigma}^2$.

Like Parker & Julliard (2003, 2005) and MMVJ (2008), we estimate models where we allow for a constant (α) in Eq. (5), even if it theoretically should not be there; in Eq. (3) there is no constant. It turns out that the constant is to play an important role for the results we find. The reason why a constant is often included in the estimation, even if it should not be there in theory, is that the constant enables the empirical model to price the cross-section of the assets the best while at the same time allowing for a constant common over- or underpricing. In other words, the constant allows the model to price the cross-section of assets the best without the additional challenge of fitting the level (the equity premium) of the returns on the assets.

We estimate Eq. (5) using GMM. The GMM procedure finds the values of the parameters γ, α, and $\mu_{c,S}$ that best satisfies the N + 1 unconditional moment conditions $E\left[h\left(\gamma, \alpha, \mu_{c,S}; \Theta_{t+1}\right)\right] = 0$ by minimizing a quadratic form of the pricing errors. When minimizing the quadratic form, we use a prespecified weighting matrix:

$$W = \begin{bmatrix} I_N & 0 \\ 0 & h \end{bmatrix}$$

like in Parker & Julliard (2005) and MMVJ (2008), such that the portfolios are given equal weight in the minimization.[6] We also follow MMVJ and set $\beta = 0.95^{1/4}$, i.e. a discount factor of five percent per annum. Other reasonable choices of β only produce negligible differences in the results.

We present two measures of fit for each of our estimations: the cross-sectional R^2 and the Hansen-Jagannathan (1997) distance. Furthermore, we present p-values from tests of whether the HJ-distances are statistically distinguishable from zero.[7]

[6] We set h to a large value in order to pin down the factor mean exactly.

[7] For further details on the computation of the HJ distance and the statistical tests of whether it is

3 Data

In this section, we briefly discuss the data we use. Summary statistics are provided in Table 1. Further details on data definitions, sources and dataset construction are provided in Appendix B.

Consumption Data. In order to investigate the long-run consumption CAPM in an international asset pricing context, the natural choice of consumption is world consumption growth. Due to data availability, we limit ourselves to quarterly private total consumption time series of the G-7 countries, taken from the IMF/IFS database. The countries' consumption growth rates (real, per capita) are weighted according to the individual country's share in real G-7 GDP.[8]

Test Assets. The vast majority of papers in the international asset pricing literature employs returns on international equity indices to conduct empirical tests of international asset pricing models (e.g. Dumas & Solnik, 1995 and DeSantis & Gerard, 1997). We follow this common practice and use returns on aggregate G-7 stock market indices (including reinvested dividends). The country indices (expressed in U.S. Dollars) are taken from Datastream and cover the period 1973Q2-2005Q4.

Our second set of equity portfolios are international book-to-market sorted portfolios. It is well-known that value stocks have historically offered a higher return than growth stocks, which has been difficult to rationalize by traditional asset pricing models (e.g. Fama and French 1993). Fama and French (1998) document similar patterns for international stock markets. The data we use are taken from Kenneth French's website. For the U.S., we use 6 size and book-to-market sorted portfolios. For France, Germany, Japan and the United Kingdom we use the international Fama/French value (high book-to-market) and growth portfolios.[9] For each non-U.S. market there is both a high book-to-market

equal to zero, see Jagannathan & Wang (1996) and Parker & Julliard (2005).

[8] This is a common procedure which has also been used in other papers exploring international aspects of consumption-based asset pricing models, such as Harvey (1989), Sarkissian (2003), or Li and Zhong (2005) among others. For robustness, we also estimate some of the models based on US consumption of non-durables and services, which is the series commonly used in the national asset pricing literature.

[9] Canada is omitted since the time series do not cover the full sample period in the Fama/French dataset.

(value) and a low book-to-market (growth) portfolio available. Thus, in total we use 16 portfolios (expressed in U.S. Dollars) which cover the sample period 1975Q1-2005Q4.

We also investigate the performance of long-run risk models to explain international bond returns. Our data are total returns on Merrill Lynch Government Bond Indices for the G-7 (ex Italy) which are taken from Datastream.[10] For each country, we use four maturity categories: 1-3, 3-5, 5-7, and 7-10 years. Hence, in total there are 24 portfolios at the quarterly frequency, and the sample period is 1986Q2-2005Q4.

We follow the extant literature by assuming a U.S. representative investor, i.e. all returns are expressed in U.S. Dollars. We subtract the U.S. 3-month T-Bill rate from the returns on the test assets when computing excess returns.

In order to investigate risk premia in the foreign exchange (FX) market, we also form currency portfolios based on international interest rate differentials. These currency portfolios thus capture excess returns related to carry trade strategies and are well known in the literature (e.g. Lustig and Verdelhan, 2007; Burnside et al. 2008; Lustig et al., 2008b). Our sample comprises currency spot (s) and 1-month forward (f) rates for 15 developed countries from November 1983 to June 2008 obtained from Datastream.[11] All exchange rates are against the USD so that we study the FX market from the perspective of a U.S. investor.

We follow the procedure of Lustig & Verdelhan (2007) when calculating the returns to carry trades. At the end of each period t, we sort currencies according to their forward discount $f_t - s_t$ and form five portfolios based on this sort. Portfolio 1 (P1) contains the 20% of currencies with the smallest forward discount (or lowest interest rate) and portfolio 5 (P5) contains the 20% of all currencies with the largest forward discount (or highest interest rates). Excess returns for portfolio j in month $t+1$ are given by $rx_{t+1}^j = f_t - s_{t+1}$ and we re-balance portfolios at the end of each month. More details regarding currency portfolios can be found in the appendix.

[10] Italy is omitted, since the time series does not cover the entire time period.

[11] These 15 countries are the same as in Lustig et al. (2008a): Australia, Austria, Belgium, Canada, Denmark, Euro area, France, Germany, Italy, Japan, Netherlands, Norway, Sweden, Switzerland, United Kingdom.

Summary Statistics. Table 1 contains descriptive statistics for our test assets' excess returns. The table contains means, standard deviations, Sharpe ratios, the minimum, and the maximum of the test assets' excess return which are all expressed in annualized percentage points. Furthermore, the statistics reported in the table include skewness, excess kurtosis, the autocorrelation coefficient of first order as well as the co-skewness measure introduced by Harvey & Siddique (2000).[12]

As shown in Panel A, annualized average excess returns in international stock markets range from about 5.6% (Japan) to about 9.0% (UK), while Sharpe ratios range from about 0.23 (Japan) to about 0.40 in the case of the U.S. Furthermore, Panel B shows that there is a clear tendency of value stocks earning higher returns than growth stocks in international equity markets; Sharpe ratios of value stocks also tend to be higher than those of growth stocks. Apart from the U.S. portfolios, there are no apparent differences of value and growth stocks in terms of their co-skewness measure in international stock markets.

Panel C reports summary statistics for the excess returns on our government bond portfolios. Typically, the average bond excess return as well as the standard deviations tend to rise with the maturity of the bonds.[13] Panel D reports summary statistics for our FX portfolios. The patterns are consistent with those reported elsewhere in the literature on the profitability of carry trade strategies (e.g. Burnside et al. 2008, Lustig and Verdelhan, 2007, Lustig et al., 2008). Contrary to what one would expect under uncovered interest rate parity (UIP), currencies with low forward discounts – i.e. low interest rates relative to the U.S. – give rise to low returns to currency speculation, while those with high forward discounts give rise to high returns. Thus, the carry trade – borrowing in the low interest rate currencies and investing in high interest rate currencies vis-a-vis the U.S. Dollar – has proved to be a very profitable strategy in the past. Likewise, Sharpe ratios differ substantially between the different forward-discount sorted portfolios.

[12] The market portfolio for computing the Harvey & Siddique (2000) measure is the excess return on the MSCI world.

[13] The overall patterns are similar to results reported elsewhere in the literature, such as Driessen et al. (2003) who – like us – also used the Merill Lynch bond indices.

4 Results

We now turn to the main results of this paper: Table 2 shows results from the estimation of Eq. (5) *without* a constant in the moment condition, whereas Table 3 shows results allowing for a constant.

Each table contains the results for the four different sets of test assets: The international stock markets in Panel A, the international value/growth portfolios in Panel B, the international bond portfolios in Panel C, and the FX portfolios in Panel D. In each panel the estimates of the risk aversion coefficient are presented, as are their associated t-statistics from tests of the hypothesis that the risk aversion coefficient is statistically distinguishable from zero. We also present the cross-sectional R^2s, the HJ-distances, and their associated p-values from tests of the hypotheses that the HJ-distances are equal to zero. In addition, the estimates of the constant are shown in Table 3.

No constant. First, we consider the case in which there is no constant in the empirical moment function, i.e. the case in which we impose the restriction that $\alpha = 0$ in Eq. (5). To explain the results, consider the results for the International Stock Markets in Panel A, Table 2. When $S = 1$, the risk-aversion coefficient is estimated to be very high: 117.84. Such high estimates of γ are well-known from the equity premium puzzle literature documented on U.S. data: A high risk-aversion coefficient is necessary to reconcile the much higher return on stocks compared to the risk-free asset with the risk of *one-period* changes in consumption.

What happens if S is larger than one? When the horizon over which consumption growth is measured is increased, i.e. S is increased, the estimate of the risk aversion coefficient is lower. This is the main point we would like to emphasize in Table 2. For instance, for $S = 2$, the estimate is reduced to 80.08, when $S = 4$, $\widehat{\gamma} = 45.62$ etc. The lowest value for the estimate of the risk aversion coefficient is obtained when $S = 12$ ($\widehat{\gamma} = 28.97$). The pattern of a lower estimate of the risk aversion coefficient in the models with no constant is common to all sets of tests assets: For the value/growth portfolios in Panel B, $\widehat{\gamma} = 334.36$ when $S = 1$, but $\widehat{\gamma} = 52.11$ when $S = 12$; for the international bond portfolios, $\widehat{\gamma} = 360.76$ when $S = 1$, but $\widehat{\gamma} = 84.12$ when $S = 12$; finally for the FX portfolios , $\widehat{\gamma} = -467.68$ when $S = 1$, but $\widehat{\gamma} = 32.75$ when $S = 12$. Hence, we find that the estimate of the risk-aversion coefficient is more reasonable when S is higher, and when there is no constant in the empirical moment function.

We do not find, however, that the model prices the cross-country distribution of returns better when S is higher. Indeed, the R^2s are all very low (and most of them even negative) regardless of the number of future consumption growth rates that are included in the model. Likewise, the HJ-distance is not lower the higher S is.[14] In other words, when there is no constant in the empirical moment function, long-run risk does not help capture differences in average returns across international assets.

A constant. What happens if there is a constant in the empirical moment function? Table 3 provides the answer: The constant generates two effects. First, the constant eliminates the tendency of a lower estimate of the risk-aversion coefficient the higher S is. Second, the constant allows for a better fit of the cross-sectional distribution of returns, the higher S is.

To be more precise, and again using Panel A as an example, the risk aversion coefficient is estimated to be 13.26 when $S = 1$, but is estimated to be negative when $S = 2$, positive when $S = 4$, and negative with $S = 8$, etc. There is thus simply no clear pattern in the estimate of γ. The same finding of no clear relation between the value of S and the estimate of the risk aversion coefficient is visible in the Panels B, C, and D. The reason why we find that the estimates of the risk-aversion coefficients behave "strangely" when changing S when there is a constant in the empirical moment function is that there is not enough cross-sectional variation in the covariances between returns and consumption growth rates (the consumption betas); The consumption betas are all more or less the same, also for larger values of S. Consequently, the betas act as a kind of "second constants" in the models, thereby creating multicollinearity-like problems. We investigate more directly the patterns of the consumption betas in the next section.

On the other hand, in Panel A the R^2 is estimated to be 0 for $S = 1$, 2, and 4, but

[14] The HJ-distance measures the largest pricing error generated by the stochastic discount factor per unit of the payoff norm. In other words, when the HJ-distance is 0.12, for instance, as in Panel A for $S = 1$, the stochastic discount factor prices the assets with a maximum pricing error of 12% compared to the observed market prices.

consequently increases to 0.06 for $S = 8$, to 0.25 for $S = 12$, and reaches 0.47 for $S = 20$. In other words, allowing for long-run risk helps capture the cross-country variation in international equity returns, when there is a constant included in the empirical moment function. Long-run risk also helps on capturing the variation in the international bond portfolios, but only does so to a lesser extent in the international value/growth portfolios and the FX portfolios.

PJ and MMVJ both use a constant in the empirical moment function. PJ and MMVJ find, for the U.S. 25 Fama & French portfolios, that the risk aversion coefficient is estimated to be lower and the cross-sectional variation in returns is captured better, when S is higher. Using international data, we can not verify that the risk aversion coefficient is lower when S is higher (and there is a constant in the empirical moment function), but we can verify that the cross-sectional fit is better.

We can visualize the better cross-sectional fit. In figures 1 and 2, we show cross-plots of the realized average excess returns along the horizontal axis and the fitted average excess returns along the vertical axis for the models with a constant in the empirical moment function, i.e. for the models of Table 3. We show the plots for $S = 1$ and $S = 8$. As is clear, the average returns are more spread out along the 45-degree line when $S = 8$ than when $S = 1$ for the international equity indices and the international bond portfolios, whereas this is not the case for the value/growth and the FX portfolios.

5 What drives these results?

We report a lower value of the risk-aversion coefficient when S is high and there is no constant in the empirical moment function. We also report that the cross-sectional R^2 is generally higher when S is high and there is a constant in the empirical moment function. Hence, regardless whether the constant is included or not, we find more reasonable results for the long-run consumption-based model along one of the two dimensions: the level of the estimate of γ or the cross-sectional R^2. On the other hand, it is also clear that the results we report for international assets are not as favorable to the long-run risk story as the results from the studies using U.S. data are. The studies using U.S. data generally include a constant in the empirical moment function, but report both lower estimates of

γ and higher cross-sectional R^2s when S is high.[15] What could be the reason for the fact that our results are not so clear-cut?

The major reason why Parker (2001, 2003), PJ (2003, 2005), and MMVJ (2008) report that long-run consumption risk plays a role in the pricing of assets cross-sectionally is that the correlation between asset returns and consumption growth increases with S and it increases in the right way for the right assets.[16] In the same vein, one reason why we do not find such a robust role for long-run risk could be that our covariances between returns and consumption growth rates do not increase in a systematic way with S.

One way to visualize this is provided in Figure 3. In part (a) this figure shows the covariances between one-period excess returns from the different stock markets and consumption growth for different values of S. As is clear from the figure, there is no clear tendency for the covariances to be higher for a higher value of S. While there are some countries where this does seem to be the case, for instance in Japan, there is generally no clear tendency.[17] The same pattern applies for the international bond portfolios (part e) and the international FX portfolios (part f). For the international value/growth portfolios (part b), there is a clear increase in the correlations when increasing S from $S = 1$ to $S = 2$. However, after $S = 2$, there is no clear increase in the correlations.

Parker & Julliard (2003, 2005) and MMVJ (2008) report that the covariances between the returns on the 25 Fama & French portfolios and consumption growth increase with S. In part (c), we "zoom in" on the U.S. value/growth portfolios. We visualize here why it is easier to capture the cross-sectional differences in the returns on the U.S. value/growth portfolios when using long-run consumption risk as a risk factor; in the next section, we formally estimate the model using U.S. value and growth portfolios only.

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[15] In their Table IV, Panel A, MMVJ report results from estimations where the constant is restricted to zero. They do not provide the cross-sectional R^2, though.

[16] For instance, Parker & Julliard (2005), page 202, write: "Because the contemporaneous covariance between returns and consumption growth is so small, a small amount of predictability, in the right pattern across assets, leads to a large increase in the relationship between consumption risk and expected returns with S."

[17] The Japanese case illustrates why the model has a hard time when confronted with international data. As Table 1 showed it is the Japanese stock market that has yielded the lowest average return over the sample period. In other words, the country with the highest consumption risk is the country with the lowest return – the opposite of what should be expected.

We can also provide more formal statistical evidence that the covariances between returns and consumption growth do not change in a strong way when increasing the horizon over which consumption growth is measured. In order to do so, Table 4 presents results from regressions of the form:

$$\sum_{s=0}^{S-1} \beta^s \Delta c_{t+1+s} = \alpha + b_1 PC1_t + b_2 PC2_t + \varepsilon_{t+1+s} \tag{6}$$

in the columns on the left-hand-side (I.A, II.A, III.A, and IV.A) and from:

$$\sum_{s=0}^{S-1} \beta^s \Delta c_{t+2+s} = \alpha + b_1 PC1_t + b_2 PC2_t + \varepsilon_{t+2+s} \tag{7}$$

in I.B, II.B, III.B, and IV.B. The difference between the two regressions is that the contemporaneous consumption growth rate is not included in the sum of discounted growth rates in Eq. (7).[18] The explanatory variables we use are the first two principal components from the set of excess returns.[19] Results based on more than the first two principal components show patterns similar to those reported in Table 4.

Consider the left-hand side of Table 4 first, i.e. I.A, II.A, III.A, and IV.A. One can witness that when increasing S from $S = 1$ to $S = 2$, i.e. including one future growth rate in consumption in addition to the contemporaneous growth rate of consumption, one of the principal components in part II.A becomes significant (and the R^2 increases, too). For $S = 4$, 8, and 12, there is no further increase in the R^2, though. For the other test assets, long-run risk does not increase the number of significant variables nor the R^2s. This conclusion is reaffirmed when excluding the contemporaneous effect on consumption, as shown in the right-hand side of the table.[20] In summary, there is only scant evidence of a stronger predictability when S is larger.

The results in Table 4 shed light on why the results for our international dataset differ along some dimensions from the results reported for the U.S. For the U.S., Parker & Julliard (2005) and MMVJ (2008) report that correlations between returns and consumption

[18] Cochrane (2007), page 284, explicitly asks for regressions such as those in Eq. (7), such that the forecastability of future growth rates can be evaluated.

[19] The first two principal components explain between 67% (International value/growth portfolios) and 88% (international FX portfolios) of the variance of the returns.

[20] In an appendix that is available on our webpages, we provide covariances between individual series of returns and the world consumption growth rate series for increasing values of S. We also provide t-statistics for these individual covariances being statistically different from zero.

growth increase when S is higher. In addition, they show that this increase in correlations is related to the cross-sectional variation in returns. For this reason, they find that cross-sectional R^2s increase, and estimates of the risk-aversion parameter fall, the higher S is. We show that even if consumption growth is not completely unpredictable (there are some significant coefficients in Table 4), the degree of predictability does not increase with S for $S > 2$, with the possible exception of the international stock market portfolios. Given that there are no clear increases in the correlations, it is not surprising that we find no clear and consistent pattern in the estimates of the risk-aversion coefficient.

6 Robustness

In this section, we evaluate whether the results presented above are robust along different dimensions.

U.S. data. We use a world-consumption series as our measure of the consumption risk factor. In order to evaluate whether our results differ in some dimensions from those reported in PJ and MMVJ since we use another (non-U.S.) consumption series, we estimate models using U.S. consumption (of nondurables and services) and compare them with the results we get when we use the world consumption series. In order to make the comparison between the results most clear, we use a standard set of U.S. test assets: the 25 Fama & French portfolios. We show the results in Table 5.

Starting from the bottom of the table (Panel D), we first show that when conducting the "standard" regression with U.S. consumption of nondurables and services as the consumption factor, the 25 Fama & French portfolios as test assets, and a constant included in the pricing equation, we find results comparable to those of Parker & Julliard (2005) and MMVJ (2008). We find an estimate of $\gamma = 120.20$ and a cross-sectional R^2 of 0.12 when $S = 1$ and an estimate of $\gamma = 79.51$ and a cross-sectional R^2 of 0.42 when $S = 12$. i.e. a reduction of the estimate of γ with around 33% and a large increase in the cross-sectional R^2.[21] We also find that the results in Table 5 are not as dependent upon the

[21] As a comparison, MMVJ find an estimate of $\gamma = 97.01$ and a cross-sectional R^2 of 0.18 when $S = 1$ and an estimate of $\gamma = 51.85$ and a cross-sectional R^2 of 0.64 when $S = 12$ (Table II, Panel D, NIPA data), i.e. a reduction of the estimate of γ by 47%. In other words, the patterns and magnitudes that we

inclusion of a constant in the pricing equation, as the results for the international data are. For instance, we find high and positive cross-sectional R^2s in Panel C (where there is no constant) using long-horizon consumption growth, i.e. when $S = 8$, 12, and 20, whereas we generally find negative R^2s in Table 2 that showed results for the international data.

Finally, regarding the use of the world consumption series, we can compare Panel A with Panel C, and Panel B with Panel D. We see that, qualitatively, there is not much difference between the results. Indeed, in both Panels A and C the estimates of γ are very high when $S = 1$ but are clearly reduced when using long-run consumption risk, up to $S = 4$.

Comparison with Parker & Julliard. We have also estimated our models using the approach of Parker & Julliard (2005), i.e. estimated Eq. (4). Whereas Eq. (5) is based on a linearization of the Euler equation, Eq. (4) is exact. On the other hand, Eq. (4) is based on the power utility function, which implies certain undesirable features, such as the restrictive assumption $\gamma = 1/\rho$.

We show the results of these tests in Table 6. Qualitatively, we find the same results as those reported above when we use the MMVJ approach: The estimates of the risk-aversion parameter decline in value the higher is S, when there is no constant in the empirical moment function, but the R^2s do not increase. On the other hand, when there is a constant in the empirical moment function, there is no clear pattern in the estimates of the risk-aversion coefficients, but the R^2s generally increase with S.

One small notable difference to the previous results, though, is that the estimates of the risk-aversion parameter are more precise in Table 6 when there is no constant in the empirical moment function (the γs are all significantly different from zero, with the exception of the estimates based upon the FX portfolios), whereas this was less so in Table 2.

Excluding Japan. As mentioned above, the Japanese stock market has performed the worst in our sample of stock markets. At the same time, the covariances of the

find are similar to those of MMVJ (the differences that do exist between our results and theirs are most likely due to the longer sample period used in MMVJ).

Japanese stock market with long-run risk increases the most with S. Thus, one may ask the question whether the results we report are in connection with the Japanese stock market behaving in a "strange" way. In Table 7, we verify that this is not the case. Indeed, in Table 7 we show results from estimations where we have excluded Japan.[22] The overall conclusions hardly change – even when the results for the international bond portfolios appear somewhat more "agreeable" (in the sense that the estimates of γ are considerably lower for $S > 2$, and the R^2s are relatively high).

Estimating since the late-1980s. Bekaert, Harvey, Lundblad & Siegel (2008) find that ever since the late-1980s, international developed equity markets are not more segmented than U.S. equity markets have been since the late 1980s. For this reason, we estimate our model for this more recent sub-period. The results are available in an appendix on our websites. We find that when we exclude the constant from the moment function, the estimates of γ are more reasonable when S is high. We also find that the cross-country R^2s are low or negative. On the other hand, when there is a constant in the moment function, the estimates of γ are imprecise, while the R^2s are reasonably high. All in all, we find the same kind of results when estimating the models over the last two decades as we do when estimating them over the full sample period.

7 Conclusion

Results in the literature based upon U.S. data show that incorporating long-run risk generally brings down estimates of the risk-aversion parameter to more reasonable levels. In addition, long-run risk often implies that the cross-sectional fit increases when using U.S. data.

We have estimated long-run consumption based asset pricing models using international tests assets. We have found that long-run risk generally either reduces estimates of risk-aversion coefficients or increases the cross-country fit, but not both. When the constant is not included – as it should be the case from a theoretical perspective – the estimate of the risk-aversion coefficient is lower when the number of periods of future consumption growth that is included in the stochastic discount factor is increased. This corresponds to

[22] We did not form new FX portfolios excluding Japan. For this reason, there is no Panel D in Table 7.

the results based upon U.S. data, however, unlike the U.S. results the explanatory power for the cross-sectional dispersion of returns is not improved. On the other hand, when a constant is included in the empirical moment function, the cross-sectional dispersion in returns is captured better when taking into account long-run risk. The estimate of the risk-aversion coefficient is not improved, though.

We conclude that the LLR-CAPMs that have been estimated on U.S. data work along some, but not all, dimensions when confronted with international data. This raises the question of how the models can be improved in order for them to perform just as well as the models developed for the U.S. equity market. Possible extensions that could help reach this goal could be estimations of models that allow for an elasticity of intertemporal substitution that is different from one. One could also consider allowing for time-variation in the volatility of consumption growth, as in the original Bansal & Yaron (2004) paper. Finally, one could include instruments in the GMM estimation. All these extensions would make the model more flexible and thus more likely to fit returns better. We leave these exciting extensions to future research.

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Appendix A: Theoretical derivations

This is a self-contained appendix providing details on the asset pricing moment conditions in the long-run consumption risk framework of MMVJ (2008) and their empirical implementation in this paper. The theoretical setup of MMVJ (2008) is based on a recursive specification of utility following Epstein & Zin (1989) as given by Equation (1) of the main text.

One major implication of the Epstein & Zin framework is the separation of risk aversion and intertemporal substitution ($\gamma \neq \frac{1}{\rho}$). The model nests the standard power utility as a special case if $\gamma = \frac{1}{\rho}$. With utility defined in Eq. (1), the stochastic discount factor (SDF) is given by

$$M_{t+1} = \beta \left(\frac{C_{t+1}}{C_t} \right)^{-\frac{1}{\rho}} \left[\frac{V_{t+1}}{E_t[V_{t+1}^{1-\gamma}]^{\frac{1}{1-\gamma}}} \right]^{\frac{1}{\rho}-\gamma}. \tag{8}$$

As Eq. (8) shows, there is an additional part in the case where risk aversion differs from the inverse of the EIS beside the conventional contemporaneous consumption growth term of the power utility model. This second term captures risk due to shocks to the investor's future consumption prospects (via the utility index).

As mentioned in the text – instead of modelling consumption growth as i.i.d. – the consumption dynamics assumed in the long-run consumption risk literature (e.g. HHL 2008, MMVJ 2008) exhibit a small predictable component. The consumption dynamics given by Equation (2) can be re-expressed as a MA(∞)-process for consumption growth

$$\begin{aligned} \Delta c_t &= \mu_c + \lambda(L)\omega_t \\ &= \mu_c + \sum_{s=0}^{\infty} \lambda_s \omega_{t-s}, \end{aligned} \tag{9}$$

where $\lambda(L)$ denotes a Lag-polynomial. HHL (2008) and MMVJ (2008) put their main focus on the asset pricing implications of the model for the case of an EIS equal to one.[23] In the $\rho = 1$ case, the log-SDF can be conveniently approximated (also see Cochrane 2007, Appendix) and tractable expressions for the stochastic discount factor can be obtained. As shown by MMVJ (2008), the assumption of $\rho = 1$ together with the assumed consumption

[23] MMVJ provide several arguments as to why other values of the EIS do not have a large impact on cross-sectional asset pricing tests. This is also confirmed in their empirical tests.

dynamics yields the following expression for the (log) SDF m_{t+1}:

$$\begin{aligned} m_{t+1} &= \ln\beta - [\mu_c + \lambda(L)\omega_{t+1}] + (1-\gamma)\lambda(\beta)\omega_{t+1} - 0.5(1-\gamma)^2\lambda(\beta)^2 \\ &= \ln\beta - \Delta c_{t+1} + (1-\gamma)\left(\sum_{s=0}^{\infty}\lambda_s\beta_s\right)\omega_{t+1} - 0.5(1-\gamma)^2\sum_{s=0}^{\infty}\lambda_s\beta_s^2 \\ &\simeq \ln\beta - (1-\gamma)\left[(E_{t+1}-E_t)\sum_{s=0}^{\infty}\beta_s\Delta c_{t+1+s}\right] - 0.5(1-\gamma)^2\sum_{s=0}^{\infty}\lambda_s\beta_s^2. \end{aligned} \tag{10}$$

The basic asset pricing equation for a risky asset i is characterized by the moment condition $E_t(M_{t+1}R^i_{t+1}) = 1$. Given the expressions for the (log) SDF m_{t+1} in (10) and assuming lognormality of the consumption growth and returns, the log-linearized Euler equation for the excess return on asset i takes the following form:

$$\begin{aligned} E_t(r^{ei}_{t+1}) + 0.5\sigma^i_t &= -Cov_t\left[m_{t+1}, r^{ei}_{t+1}\right] \\ &\simeq (\gamma-1)Cov_t\left[(E_{t+1}-E_t)\sum_{s=0}^{\infty}\beta_s\Delta c_{t+1+s}, r^{ei}_{t+1}\right] \\ &= (\gamma-1)Cov_t\left[\sum_{s=0}^{\infty}\beta_s\Delta c_{t+1+s}, r^{ei}_{t+1}\right] \\ &\quad -(\gamma-1)Cov_t\left[E_t\sum_{s=0}^{\infty}\beta_s\Delta c_{t+1+s}, E_t(r^{ei}_{t+1})\right]. \end{aligned} \tag{11}$$

Empirical Implementation The empirical asset pricing tests in this paper are based on unconditional versions of the moment conditions defined through Eq. (11). Our implementation of the long-run risk framework closely follows MMVJ in that we estimate the model with the unconditional covariance term (i.e. the first term) in Eq. (11), whereas the covariance of the conditional expectation of the present value of consumption growth and conditionally expected returns is neglected.[24] The discount factor β is set to 5% at the annual level which implies $0.95^{0.25}$ for the quarterly data.

Appendix B: Data

This appendix describes the data series we use and their sources.

[24] Similar to MMVJ, our international asset return data do not span a very long-term time period compared to long-term US data. Thus, alternative approaches obtaining a measure of conditional expectations through a VAR may induce a substantial amount of measurement error.

CONSUMPTION: G-7 consumption data (total private consumption) are taken from the IMF/IFS database (accessed via Datastream) and deflated by the respective Consumer Price Index CPI (also taken from the IMF/IFS). The country's consumption series are converted into U.S. dollars and expressed in per capita terms by dividing them by population figures (IMF/IFS). In order to construct our measure of world consumption growth, the individual country consumption growth rates are aggregated with weights determined by the respective country's share in aggregate (G-7) real GDP. In Table 5, we also make use of U.S. consumption growth of non-durables and services (real, per capita). These data are taken from the NIPA tables by the Bureau of Economic Analysis (BEA).

INT. EQUITY INDEX RETURNS: We use aggregate stock market indices (including reinvested dividends) for the G-7 countries taken from Datastream. The quarterly returns are expressed in U.S. Dollars. Excess returns are computed by subtracting the U.S. 3-month T-Bill rate. The sample period is 1973Q2-2005Q4.

INT. VALUE/GROWTH PORTFOLIOS: Our source of international value/growth portfolios is the Data Library on Ken French's webpage. For the U.S., we use 6 size and book-to-market sorted portfolios, which form the basis for the Fama-French mimicking factors SMB and HML. The data for the remaining countries (France, Germany, Japan and the United Kingdom) is also taken from Kenneth French's webpage. Canada is omitted because the Canadian time series does not span the entire sample period of the international Fama/French dataset. There is both a growth (low book-to-market) and a value (high book-to-market) portfolio for each stock market available. These portfolios are expressed in U.S. Dollars. Excess returns are computed by subtracting the U.S. 3 month T-Bill rate. The sample period is 1975Q1-2005Q4.

INT. BOND RETURNS: Our international bond returns are the total returns on Merrill Lynch Government Bond Indices for the G-7 (ex Italy). The data are taken from Datastream. Italy is omitted, since the time series does not cover the entire time period. For each country, we use four bond maturity groups: 1-3, 3-5, 5-7, and 7-10 years. The bond returns are expressed in U.S. Dollars; to obtain excess returns, the U.S. 3-month T-Bill rate is subtracted. The sample period covers 1986Q2-2005Q4.

INT. FX PORTFOLIOS: Our spot and forward rate data come from BBI except for Belgium (WMF/Reuters). Only a limited amount of currencies is available for November 1983 to June 2008 so that our sample size varies. For example, several European currencies are replaced by the Euro at the start of 1999 due to the monetary union. We follow the recent literature (Lustig, Roussanov & Verdelhan 2008) by using spot and forward rates to construct currency excess returns instead of using genuine interest rate data. Using forward rates instead of interest rates appears advantageous since high-quality forward rate data is readily available on a high frequency. In contrast, international data on very short-term interest rates is not easily available and is often not comparable across countries. Using forward rates or interest rates leads to the same results, however, as the following lines show.

The log excess return of investing in foreign currency is simply given by:

$$rx_{t+1}^{j} \approx i_t^* - i_t - \triangle s_{t+1}$$

so that the log spot rate change is adjusted for differences in interest rates paid at home (i) and earned abroad (i^*). Since covered interest parity holds (see e.g. Akram et al. 2008), the forward discount $f - s$ has to satisfy

$$f_t - s_t \approx i_t^* - i_t$$

and sorting on forward discounts is equivalent to sorting on interest rate differentials. Inserting the last into the former equation finally results in

$$rx_{t+1}^{j} \approx f_t - s_t - \triangle s_{t+1} = f_t - s_{t+1}.$$

Log currency excess return $rx_{j,t+1}$ for portfolio j by taking the average of the log currency excess returns in each portfolio j.

Table 1: Descriptive Statistics of Portfolio Returns

Panel A: International Equity Portfolios (1973Q2-2005Q4)									
Portfolio	MEAN	STD	SR	Min	Max	SKW	KURT	COSK	AC1
Canada	5.858	18.304	0.320	-102.075	116.022	0.084	0.016	-0.075	0.011
France	9.624	24.744	0.389	-103.967	179.635	0.403	1.084	0.060	0.187
Germany	6.262	20.565	0.305	-82.175	142.391	0.376	0.507	-0.195	0.094
Italy	7.508	28.738	0.261	-134.623	281.061	1.106	3.797	0.079	0.149
Japan	5.594	24.202	0.231	-112.385	162.392	0.526	0.869	0.195	0.210
United Kingdom	9.003	23.527	0.383	-107.171	188.557	0.526	1.693	-0.103	-0.063
United States	5.829	14.374	0.406	-82.564	99.084	0.024	0.751	-0.174	0.066

Panel B: International Value/Growth Portfolios (1975Q1-2005Q4)									
Portfolio	MEAN	STD	SR	Min	Max	SKW	KURT	COSK	AC1
France (high BM)	15.420	29.629	0.520	-161.323	206.218	0.232	1.036	-0.127	-0.010
France (low BM)	8.503	24.589	0.346	-147.659	166.455	0.253	0.880	-0.173	-0.002
Germany (high BM)	13.276	24.402	0.544	-142.365	212.852	0.444	2.154	-0.279	0.021
Germany (low BM)	7.528	23.365	0.322	-154.492	128.220	-0.239	0.898	-0.235	0.033
Italy (high BM)	7.871	33.176	0.237	-136.429	327.657	1.283	4.464	-0.127	0.068
Italy (low BM)	8.621	29.638	0.291	-112.552	264.943	1.001	1.797	-0.076	0.080
Japan (high BM)	14.380	26.896	0.535	-112.342	172.087	0.068	-0.069	-0.010	0.058
Japan (low BM)	3.254	26.611	0.122	-145.989	178.958	0.297	0.734	0.151	0.041
United Kingdom (high BM)	13.914	26.094	0.533	-75.010	348.587	2.123	11.944	0.102	-0.087
United Kingdom (low BM)	10.715	24.620	0.435	-79.767	321.713	2.059	11.311	0.113	-0.045
United States S1B1	9.407	27.369	0.344	-130.202	167.738	0.042	0.285	0.099	-0.068
United States S1B2	15.070	20.912	0.721	-108.902	153.013	-0.170	1.160	0.039	-0.084
United States S1B3	17.072	21.882	0.780	-113.689	180.980	0.200	2.298	-0.032	-0.069
United States S2B1	7.828	18.491	0.423	-99.303	99.821	-0.178	0.214	0.177	0.007
United States S2B2	9.788	15.674	0.624	-92.350	83.359	-0.472	0.776	0.076	-0.039
United States S2B3	10.903	16.191	0.673	-86.733	124.864	-0.141	1.606	0.006	-0.027

Notes: The table reports descriptive statistics for the returns of the test assets. Panel A reports results for the excess returns of international G7 equity indices (Canada, France, Germany, Italy, Japan, UK and US). Panel B includes descriptives for international value/growth indices (France, Germany, Italy, Japan, UK) and six US size and book-to-market portfolios. Means (MEAN), standard deviations (STD), Sharpe ratios (SR), the minimim (Min), the maximum (Max) are expressed in annual terms (in %). SKW denotes Skewness, KURT excess kurtosis. Co-skewness (COSK) is computed according to Harvey & Siddique (2000). AC1 denotes the autocorrelation correlation coefficient of first order.

Table 1: cont.

Panel C: International Bond Portfolios (1986Q2-2005Q4)									
Portfolio	MEAN	STD	SR	Min	Max	SKW	KURT	COSK	AC1
Canada (1T3)	4.278	6.504	0.658	-28.673	38.262	0.061	0.060	0.247	-0.083
Canada (3T5)	5.457	7.785	0.701	-36.502	42.848	0.108	0.009	0.238	-0.138
Canada (5T7)	6.033	9.528	0.633	-48.237	53.695	0.360	0.386	0.401	-0.226
Canada (7T10)	6.560	9.767	0.672	-54.435	49.438	0.121	0.422	0.253	-0.175
France (1T3)	4.281	10.902	0.393	-45.712	48.002	-0.148	-0.691	0.518	0.019
France (3T5)	5.429	11.277	0.481	-42.819	59.503	0.043	-0.547	0.582	0.015
France (5T7)	6.160	11.401	0.540	-42.843	51.395	-0.020	-0.683	0.560	0.015
France (7T10)	6.775	12.042	0.563	-39.720	75.515	0.207	-0.264	0.580	-0.005
Germany (1T3)	3.658	11.860	0.308	-54.126	53.250	-0.071	-0.311	0.523	-0.021
Germany (3T5)	4.569	12.288	0.372	-53.785	57.214	0.043	-0.326	0.540	-0.026
Germany (5T7)	5.249	12.632	0.416	-52.817	59.802	0.123	-0.349	0.556	-0.043
Germany (7T10)	5.303	13.030	0.407	-52.900	64.120	0.159	-0.399	0.559	-0.077
Japan (1T3)	2.197	14.208	0.155	-56.234	94.175	0.842	0.771	0.582	0.033
Japan (3T5)	3.401	14.759	0.230	-54.170	97.953	0.911	1.044	0.553	0.025
Japan (5T7)	4.294	15.654	0.274	-53.017	104.947	0.927	1.172	0.531	-0.000
Japan (7T10)	5.009	16.882	0.297	-59.794	116.627	0.936	1.442	0.520	-0.058
United Kingdom (1T3)	5.408	11.385	0.475	-55.024	74.712	0.486	1.335	0.824	-0.112
United Kingdom (3T5)	6.022	12.474	0.483	-44.368	88.623	0.779	1.716	0.829	-0.140
United Kingdom (5T7)	6.762	13.580	0.498	-44.161	102.077	0.816	1.905	0.830	-0.158
United Kingdom (7T10)	7.232	14.591	0.496	-46.983	114.267	0.890	2.223	0.838	-0.169
United States (1T3)	1.693	2.078	0.815	-9.753	12.093	0.138	0.126	0.270	0.107
United States (3T5)	2.716	4.139	0.656	-20.269	22.400	0.064	0.035	0.268	0.021
United States (5T7)	3.293	5.307	0.620	-25.097	28.846	0.078	0.096	0.285	-0.006
United States (7T10)	3.676	6.549	0.561	-32.332	39.097	0.019	0.226	0.284	-0.012

Panel D: International FX Portfolios (1984Q2-2005Q4)									
Portfolio	MEAN	STD	SR	Min	Max	SKW	KURT	COSK	AC1
FXP1	-5.311	8.950	-0.593	-50.077	58.227	0.785	1.656	0.194	0.160
FXP2	-0.021	8.643	-0.002	-41.187	61.521	0.509	0.917	0.396	0.139
FXP3	2.875	9.268	0.310	-32.982	45.899	0.331	-0.437	0.136	0.151
FXP4	4.384	9.769	0.449	-50.749	56.386	0.043	0.245	0.224	0.118
FXP5	6.506	9.725	0.669	-66.130	58.608	-0.157	1.391	0.184	0.104

Notes: The table reports descriptive statistics for the returns of the test assets. Panel C reports results for the excess returns of international bond indices (Canada, France, Germany, Italy, Japan, UK and US). Panel D includes descriptives for international FX portfolios sorted according to forward-discounts. Means (MEAN), standard deviations (STD), Sharpe ratios (SR), the minimim (Min), the maximum (Max) are expressed in annual terms (in %). SKW denotes Skewness, KURT excess kurtosis. Co-skewness (COSK) is computed according to Harvey & Siddique (2000). AC1 denotes the autocorrelation correlation coefficient of first order.

Table 2: Estimation Results for the Euler Equation of the LR-CCAPM Across Horizons (Time-series Tests, Constant Excluded)

Panel A: International Stock Markets							
Horizon	1	2	4	8	12	16	20
$\hat{\gamma}$	117.84	80.08	45.62	40.01	28.97	46.34	-7.95
	(1.82)	(1.74)	(1.74)	(1.66)	(1.40)	(1.40)	(-0.37)
R^2	-0.34	-1.88	-1.97	-4.97	-3.57	-3.94	0.32
HJ-dist.	0.12	0.17	0.16	0.21	0.20	0.22	0.28
	(0.84)	(0.52)	(0.63)	(0.32)	(0.23)	(0.20)	(0.08)

Panel B: International Value/Growth Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\gamma}$	334.36	130.54	88.87	74.82	52.11	70.11	87.38
	(1.44)	(2.03)	(2.36)	(2.21)	(1.89)	(2.02)	(1.81)
R^2	-2.44	-1.48	-0.72	-1.49	-1.50	-1.27	-2.20
HJ-dist.	0.59	0.59	0.59	0.59	0.58	0.58	0.58
	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.03)

Panel C: International Bond Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\gamma}$	360.76	194.41	154.14	85.87	84.12	62.66	85.90
	(1.33)	(1.44)	(1.41)	(1.36)	(1.14)	(1.18)	(1.26)
R^2	-2.34	-2.32	-2.87	-2.37	-0.88	-0.39	-0.65
HJ-dist.	0.61	0.63	0.63	0.67	0.71	0.77	0.76
	(0.08)	(0.05)	(0.08)	(0.08)	(0.05)	(0.02)	(0.03)

Panel D: FX Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\gamma}$	-467.68	93.50	21.42	4.02	32.75	28.48	29.41
	(-1.41)	(1.01)	(0.32)	(0.08)	(0.92)	(0.82)	(0.82)
R^2	0.17	-0.01	-0.01	-0.01	-0.06	0.00	0.01
HJ-dist.	0.47	0.47	0.47	0.46	0.45	0.42	0.40
	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.04)	(0.09)

Notes: The table reports estimation results for the LR-CCAPM for various international asset portfolios. Estimation results are presented using discounted long-run world consumption growth over horizons $S = 1, 2, 4, 8, 12, 16, 20$. The estimation draws on the moment conditions implied by a cross-sectional regression of the average excess portfolio return (plus one half of the variance of the excess return) on the covariance of the log portfolio excess return with long-run discounted consumption growth. The estimation is performed using GMM with a prespecified weighting matrix. The constant is NOT included in the cross-sectional regression. The reported estimation results include an intercept α, the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey & West (1987) with $S + 1$ lags in parentheses), the cross-sectional R^2 and the HJ-distance (simulation-based p-value in parentheses). The set of international test assets includes G7 aggregate stock market indices (Panel A), equity portfolios sorted by book-to-market (Panel B), international bond portfolios (Panel C) and FX portfolios (Panel D). All returns are expressed in US Dollars.

Table 3: Estimation Results for the Euler Equation of the LR-CCAPM Across Horizons (Cross-sectional Tests, Constant Included)

Panel A: International Stock Markets							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.016	0.019	0.017	0.018	0.017	0.018	0.019
	(0.86)	(1.59)	(1.32)	(1.79)	(1.51)	(1.52)	(1.78)
$\hat{\gamma}$	13.26	-7.65	1.87	-3.57	-8.59	-14.14	-19.26
	(0.11)	(-0.14)	(0.07)	(-0.19)	(-0.50)	(-0.63)	(-0.76)
R^2	0.00	0.02	0.00	0.06	0.25	0.26	0.47
HJ-dist.	0.12	0.13	0.13	0.13	0.14	0.16	0.15
	(0.69)	(0.64)	(0.67)	(0.50)	(0.47)	(0.38)	(0.42)

Panel B: International Value/Growth Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.028	0.035	0.026	0.026	0.027	0.034	0.033
	(3.89)	(3.45)	(2.63)	(3.12)	(2.66)	(2.39)	(2.49)
$\hat{\gamma}$	-19.73	-45.44	3.70	2.23	-11.96	-22.67	-22.30
	(-0.26)	(-0.78)	(0.13)	(0.11)	(-0.67)	(-0.88)	(-0.74)
R^2	0.01	0.11	0.00	0.00	0.06	0.09	0.10
HJ-dist.	0.51	0.50	0.51	0.51	0.49	0.48	0.48
	(0.00)	(0.00)	(0.02)	(0.03)	(0.08)	(0.06)	(0.05)

Panel C: International Bond Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.013	0.012	0.013	0.011	0.008	0.006	0.007
	(2.48)	(2.63)	(2.65)	(2.37)	(2.01)	(1.80)	(1.82)
$\hat{\gamma}$	-26.05	3.26	4.67	17.77	30.25	24.88	33.61
	(-0.16)	(0.04)	(0.08)	(0.54)	(1.05)	(0.65)	(1.01)
R^2	0.01	0.00	0.00	0.15	0.37	0.26	0.41
HJ-dist.	0.54	0.56	0.56	0.58	0.61	0.62	0.64
	(0.10)	(0.11)	(0.13)	(0.16)	(0.13)	(0.10)	(0.07)

Panel D: FX Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.004	0.003	0.005	0.007	0.005	0.002	0.002
	(0.42)	(0.59)	(0.81)	(0.65)	(0.81)	(0.33)	(0.35)
$\hat{\gamma}$	-453.72	27.05	-49.25	-68.32	-16.06	15.16	17.05
	(-1.40)	(0.39)	(-0.86)	(-1.90)	(-0.69)	(0.85)	(1.35)
R^2	0.19	0.00	0.01	0.08	0.01	0.01	0.02
HJ-dist.	0.47	0.47	0.47	0.46	0.45	0.42	0.40
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.02)	(0.06)

Notes: The table reports estimation results for the LR-CCAPM for various international asset portfolios. Estimation results are presented using discounted long-run world consumption growth over horizons $S = 1, 2, 4, 8, 12, 16, 20$. The estimation draws on the moment conditions implied by a cross-sectional regression of the average excess portfolio return plus one half of the variance of the excess return on the covariance of the log portfolio excess return with long-run discounted consumption growth. The estimation is performed using GMM with a prespecified weighting matrix. The constant is included in the cross-sectional regression. The reported estimation results include the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey & West (1987) with $S+1$ lags in parentheses), the cross-sectional R^2 and the HJ-distance (simulation-based p-value in parentheses). The set of international test assets includes aggregate stock market indices (Panel A), equity portfolios sorted by book-to-market (Panel B), international bond portfolios (Panel C) and FX portfolios (Panel D). All returns are expressed in US Dollars.

Table 4: Common Factors of Portfolio Returns and Long-run World Consumption Growth

I.A.	International Stock Markets					I.B.	International Stock Markets				
Horizon	1	2	4	8	12	Horizon	1	2	4	8	12
PC1	1.987	2.733	4.565	3.045	3.616	PC1	0.751	1.654	2.185	0.323	1.466
	(3.80)	(2.84)	(3.41)	(1.16)	(1.00)		(1.14)	(1.74)	(1.52)	(0.11)	(0.39)
PC2	-0.817	-1.788	-4.667	-4.849	-3.910	PC2	-0.919	-2.101	-3.904	-3.649	-3.185
	(-1.77)	(-2.35)	(-3.05)	(-1.98)	(-1.60)		(-1.73)	(-2.36)	(-2.58)	(-1.62)	(-1.44)
R^2	10.84	11.09	18.52	6.31	3.55	R^2	3.37	7.51	8.86	2.68	1.57

II.A.	International Value/Growth Portfolios					II.B.	International Value/Growth Portfolios				
Horizon	1	2	4	8	12	Horizon	1	2	4	8	12
PC1	0.613	2.300	3.628	3.606	4.666	PC1	1.693	2.408	3.398	2.790	4.409
	(1.22)	(2.61)	(3.30)	(2.20)	(1.86)		(2.47)	(2.73)	(3.23)	(1.50)	(1.60)
PC2	-0.423	-0.531	0.401	0.640	0.975	PC2	-0.106	0.238	0.777	0.335	1.153
	(-0.85)	(-0.82)	(0.31)	(0.27)	(0.30)		(-0.25)	(0.35)	(0.61)	(0.15)	(0.35)
R^2	1.53	6.39	6.09	2.50	2.56	R^2	7.55	6.53	5.55	1.50	2.35

III.A.	International Bond Portfolios					III.B.	International Bond Portfolios				
Horizon	1	2	4	8	12	Horizon	1	2	4	8	12
PC1	0.499	0.982	0.654	2.357	2.463	PC1	0.516	-0.287	0.838	2.051	2.063
	(1.30)	(1.36)	(0.71)	(1.40)	(1.05)		(0.99)	(-0.44)	(0.96)	(1.27)	(1.13)
PC2	-0.419	-0.861	-1.312	-1.856	-0.736	PC2	-0.465	-0.615	-1.112	-0.619	0.015
	(-1.26)	(-1.74)	(-1.88)	(-1.59)	(-0.63)		(-1.18)	(-1.23)	(-1.40)	(-0.56)	(0.01)
R^2	2.89	6.17	4.06	7.38	4.14	R^2	3.36	1.76	3.79	4.00	3.04

IV.A.	FX Portfolios					IV.B.	FX Portfolios				
Horizon	1	2	4	8	12	Horizon	1	2	4	8	12
PC1	-0.061	1.010	0.115	0.462	1.365	PC1	1.100	0.816	0.262	0.941	1.201
	(-0.14)	(1.54)	(0.09)	(0.21)	(0.46)		(2.53)	(1.05)	(0.19)	(0.39)	(0.44)
PC2	-0.261	-0.417	-0.273	-0.983	-0.969	PC2	-0.167	-0.424	-0.853	-0.719	-0.151
	(-0.45)	(-0.64)	(-0.33)	(-0.83)	(-0.67)		(-0.54)	(-0.73)	(-1.17)	(-0.76)	(-0.11)
R^2	0.33	3.60	0.10	0.52	0.87	R^2	7.25	2.56	0.90	0.70	0.46

Notes: The table provides diagnostic regression results for investigating the relation between long-run consumption risk and common factors of the test asset excess returns R^e_{t+1}. Estimation results are presented using discounted long-run world consumption growth over horizons $S = 1, 2, 4, 8, 12$. Panel A presents results of a regression of discounted consumption growth $\sum_{s=0}^{S-1} \beta^s \Delta c_{t+1+s}$ on the first two principal components of the portfolio returns R^e_{t+1}. Panel B presents results of the forecasting regression $\sum_{s=0}^{S-1} \beta^s \Delta c_{t+2+s}$ regressed on the first two principal components of the portfolio returns R^e_{t+1}. Standard errors in Panel A/B are based on the adjustment by Newey & West (1987) with $S + 1$ lags.

Table 5: 25 Fama-French Portfolio: Estimation Results for the Euler Equation of the LR-CCAPM Across Horizons with US and World Consumption Growth

Panel A: 25 Fama-French: World Consumption Growth, without constant							
Horizon	1	2	4	8	12	16	20
$\hat{\gamma}$	525.80	162.13	99.84	127.64	90.09	108.09	154.81
	(1.39)	(2.03)	(2.18)	(1.66)	(2.11)	(2.26)	(1.79)
R^2	-1.54	-0.49	0.17	0.17	-0.25	-0.77	-1.59
HJ-dist.	0.61	0.61	0.62	0.62	0.62	0.60	0.57
	(0.00)	(0.00)	(0.00)	(0.01)	(0.04)	(0.17)	(0.36)

Panel B: 25 Fama-French: World Consumption Growth, with constant							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.026	0.029	0.007	0.014	0.018	0.024	0.023
	(2.81)	(2.53)	(0.42)	(0.56)	(1.04)	(1.85)	(2.42)
$\hat{\gamma}$	137.56	14.83	78.28	74.50	35.48	29.59	48.59
	(1.43)	(0.21)	(1.37)	(1.04)	(0.92)	(0.89)	(1.43)
R^2	0.23	0.00	0.18	0.36	0.16	0.12	0.40
HJ-dist.	0.56	0.55	0.56	0.57	0.59	0.57	0.54
	(0.00)	(0.00)	(0.01)	(0.01)	(0.03)	(0.10)	(0.36)

Panel C: 25 Fama-French: US Consumption Growth, without constant							
Horizon	1	2	4	8	12	16	20
$\hat{\gamma}$	366.78	146.50	122.22	108.74	107.29	151.36	165.57
	(1.92)	(2.36)	(2.11)	(1.98)	(2.12)	(2.05)	(2.08)
R^2	-0.35	-0.14	-0.06	0.62	0.37	-0.58	0.18
HJ-dist.	0.60	0.60	0.62	0.61	0.60	0.57	0.57
	(0.00)	(0.00)	(0.00)	(0.03)	(0.12)	(0.25)	(0.25)

Panel D: 25 Fama-French: US Consumption Growth, with constant							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.022	0.020	0.018	-0.003	0.008	0.021	0.015
	(2.53)	(1.88)	(1.68)	(-0.11)	(0.35)	(1.75)	(1.42)
$\hat{\gamma}$	120.20	56.89	56.22	118.07	79.51	56.55	89.65
	(1.10)	(0.97)	(1.24)	(1.29)	(1.19)	(1.37)	(1.24)
R^2	0.12	0.10	0.14	0.63	0.42	0.30	0.67
HJ-dist.	0.56	0.56	0.56	0.57	0.58	0.56	0.54
	(0.00)	(0.00)	(0.01)	(0.02)	(0.07)	(0.19)	(0.24)

Notes: The table reports estimation results for the LR-CCAPM for the Fama-French 25 Size and Book-to-Market sorted Portfolios. Estimation results are presented using discounted long-run world and US consumption growth over horizons $S = 1, 2, 4, 8, 12, 16, 20$. The estimation draws on the moment conditions implied by a cross-sectional regression of the average excess portfolio return (plus one half of the variance of the excess return) on the covariance of the log portfolio excess return with long-run discounted consumption growth. The estimation is performed using GMM with a prespecified weighting matrix. The reported estimation results include an intercept α, the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey & West (1987) with $S + 1$ lags in parentheses), the cross-sectional R^2 and the HJ-distance (simulation-based p-value in parentheses).

Table 6: Estimation Results: Parker & Julliard Specification

Cross-Sectional Tests (Const. excluded)					Cross-Sectional Tests (Const. included)				
I.A	International Stock Markets				I.B	International Stock Markets			
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	–	–	–	–	$\hat{\alpha}$	0.018	0.018	0.017	0.019
	(–)	(–)	(–)	(–)		(0.87)	(1.29)	(1.63)	(1.75)
$\hat{\gamma}$	76.29	23.59	9.74	9.14	$\hat{\gamma}$	-3.07	0.47	-11.58	-52.48
	(2.96)	(3.30)	(3.16)	(4.15)		(-0.02)	(0.02)	(-0.33)	(-0.13)
R^2	-0.41	-2.00	-3.43	-3.84	R^2	0.00	0.00	0.21	0.24
HJ-dist.	0.13	0.17	0.20	0.23	HJ-dist.	0.12	0.13	0.14	0.16
	(0.93)	(0.82)	(0.46)	(0.32)		(0.88)	(0.87)	(0.76)	(0.60)

II.A	International Value/Growth Portfolios				II.B	International Value/Growth Portfolios			
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	–	–	–	–	$\hat{\alpha}$	0.028	0.026	0.027	0.034
	(–)	(–)	(–)	(–)		(3.75)	(2.45)	(2.71)	(2.54)
$\hat{\gamma}$	130.04	31.86	12.08	10.19	$\hat{\gamma}$	-24.16	4.41	-23.58	72.06
	(3.65)	(6.75)	(5.46)	(6.01)		(-0.24)	(0.19)	(-0.32)	(0.21)
R^2	-2.38	-0.59	-1.44	-1.40	R^2	0.01	0.00	0.06	0.09
HJ-dist.	0.60	0.60	0.59	0.60	HJ-dist.	0.52	0.52	0.50	0.50
	(0.00)	(0.00)	(0.02)	(0.02)		(0.01)	(0.03)	(0.10)	(0.08)

III.A	International Bond Portfolios				III.B	International Bond Portfolios			
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	–	–	–	–	$\hat{\alpha}$	0.013	0.013	0.009	0.006
	(–)	(–)	(–)	(–)		(2.52)	(2.63)	(2.02)	(1.96)
$\hat{\gamma}$	137.76	41.27	15.52	11.36	$\hat{\gamma}$	-20.61	3.96	10.38	7.68
	(3.13)	(4.16)	(3.62)	(3.84)		(-0.11)	(0.08)	(2.08)	(1.23)
R^2	-2.25	-2.78	-1.02	-0.45	R^2	0.01	0.00	0.36	0.25
HJ-dist.	0.61	0.64	0.71	0.78	HJ-dist.	0.54	0.56	0.61	0.63
	(0.22)	(0.18)	(0.07)	(0.03)		(0.40)	(0.30)	(0.18)	(0.12)

IV.A	FX Portfolios				IV.B	FX Portfolios			
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	–	–	–	–	$\hat{\alpha}$	0.004	0.006	0.005	0.001
	(–)	(–)	(–)	(–)		(0.43)	(0.65)	(0.80)	(0.22)
$\hat{\gamma}$	398.39	-90.01	10.23	8.23	$\hat{\gamma}$	413.17	137.74	-176.45	6.79
	(1.73)	(-0.17)	(1.88)	(1.86)		(1.62)	(1.10)	(-0.09)	(2.17)
R^2	0.20	0.02	-0.06	0.01	R^2	0.20	0.04	0.01	0.02
HJ-dist.	0.47	0.47	0.45	0.42	HJ-dist.	0.47	0.47	0.45	0.42
	(0.00)	(0.00)	(0.02)	(0.05)		(0.00)	(0.00)	(0.01)	(0.03)

Notes: The table reports estimation results of Parker and Julliard's long-run CCAPM specification for various international asset portfolios. Estimation results are presented using long-run world consumption growth over horizons $S = 1, 2, 4, 8, 12, 16, 20$. The estimation is performed using GMM with a prespecified weighting matrix. The reported estimation results include an intercept α, the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey & West (1987) with $S + 1$ lags in parentheses), the cross-sectional R^2 and the HJ-distance (simulation-based p-value in parentheses). The set of international test assets includes G7 aggregate stock market indices (Panel A), equity portfolios sorted by book-to-market (Panel B), international bond portfolios (Panel C) and FX portfolios (Panel D). All returns are expressed in US Dollars.

Table 7: International Equity and Bond Portfolios: Estimation Results for the Euler Equation of the LR-CCAPM Across Horizons, Cross-sectional Tests Excluding Japan

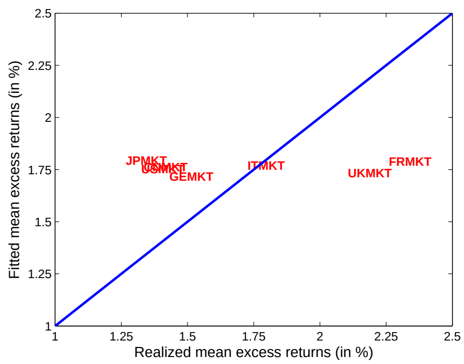
Panel A: International Stock Markets (ex Jap)							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.008	0.010	0.014	0.017	0.016	0.018	0.019
	(0.40)	(0.61)	(1.02)	(1.74)	(1.43)	(1.51)	(1.73)
$\hat{\gamma}$	73.90	47.26	14.72	3.04	-7.97	-9.52	-24.16
	(0.50)	(0.54)	(0.45)	(0.14)	(-0.44)	(-0.41)	(-0.79)
R^2	0.14	0.24	0.15	0.01	0.10	0.08	0.41
HJ-dist.	0.10	0.11	0.10	0.12	0.13	0.14	0.14
	(0.59)	(0.62)	(0.63)	(0.30)	(0.21)	(0.18)	(0.15)

Panel B: International Value/Growth Portfolios (ex Jap)							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.027	0.032	0.022	0.022	0.021	0.023	0.028
	(3.67)	(3.25)	(2.08)	(2.20)	(1.88)	(2.16)	(3.36)
$\hat{\gamma}$	4.42	-23.09	21.63	22.86	11.34	15.25	2.32
	(0.06)	(-0.39)	(0.60)	(0.74)	(0.50)	(0.63)	(0.12)
R^2	0.00	0.04	0.06	0.13	0.03	0.03	0.00
HJ-dist.	0.44	0.44	0.45	0.44	0.44	0.44	0.41
	(0.01)	(0.01)	(0.01)	(0.05)	(0.05)	(0.06)	(0.16)

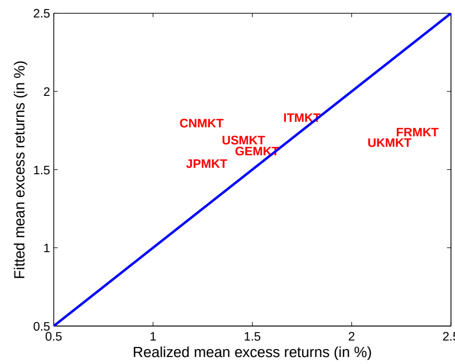
Panel C: International Bond Portfolios (ex Jap)							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.010	0.006	0.013	0.009	0.008	0.006	0.007
	(1.55)	(0.90)	(2.47)	(2.37)	(2.13)	(2.51)	(2.17)
$\hat{\gamma}$	161.40	207.38	39.82	56.37	44.66	29.67	38.69
	(0.73)	(1.65)	(0.52)	(1.35)	(1.61)	(0.82)	(1.13)
R^2	0.12	0.48	0.08	0.69	0.66	0.31	0.39
HJ-dist.	0.51	0.52	0.53	0.53	0.54	0.54	0.57
	(0.06)	(0.07)	(0.10)	(0.13)	(0.10)	(0.10)	(0.05)

Notes: The table reports estimation results for the LR-CCAPM for international stock markets (Panel A), international value/growth portfolios (Panel B) and bond portfolios (Panel C). Assets from Japan are excluded from the set of test assets. Estimation results are presented using discounted long-run world consumption growth over horizons $S = 1, 2, 4, 8, 12, 16, 20$. The estimation draws on the moment conditions implied by a cross-sectional regression of the average excess portfolio return (plus one half of the variance of the excess return) on the covariance of the log portfolio excess return with long-run discounted consumption growth. The estimation is performed using GMM with a prespecified weighting matrix. The reported estimation results include an intercept α, the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey & West (1987) with $S + 1$ lags in parentheses), the cross-sectional R^2 and the HJ-distance (simulation-based p-value in parentheses).

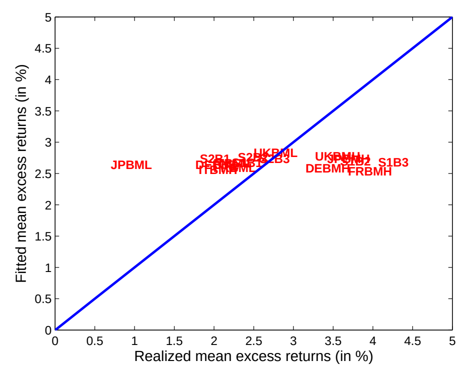
Figure 1: Pricing Errors for International Equity Portfolios



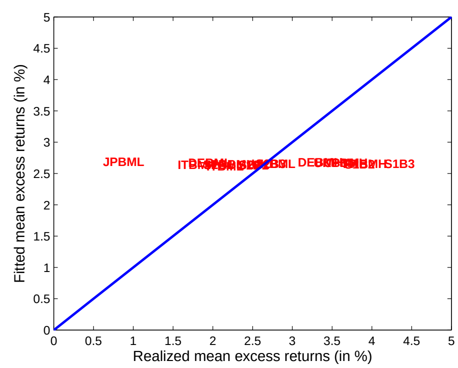
(a) Int. Equity Indices, S=1



(b) Int. Equity Indices, S=8



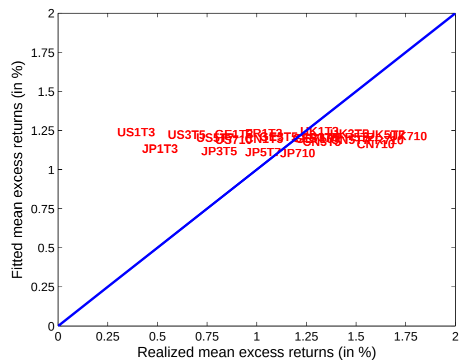
(c) Int. BM Portfolios, S=1



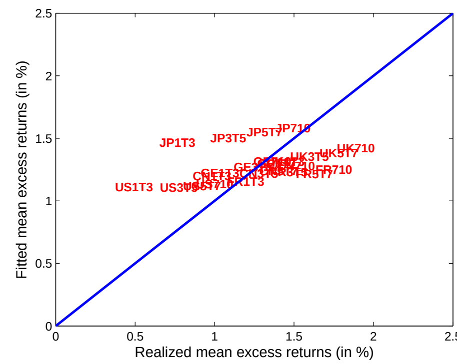
(d) Int. BM Portfolios, S=8

Notes: The Figure depicts pricing error plots for two sets of international equity portfolios for horizons of long-run consumption growth of $S = 1$ and $S = 8$. The constant is included in the cross-sectional regression. Subfigures (a) and (b) use international aggregate stock market indices as test assets, while subfigures (c) and (d) are based on international value/growth portfolios.

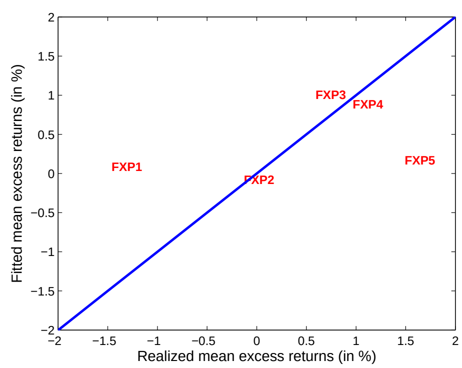
Figure 2: Pricing Errors for International Bond and FX Portfolios



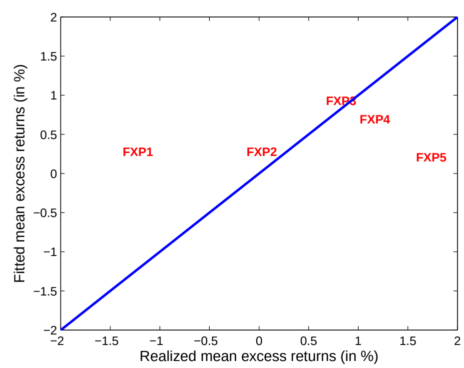
(a) Int. Bond Portfolios, S=1



(b) Int. Bond Portfolios, S=8



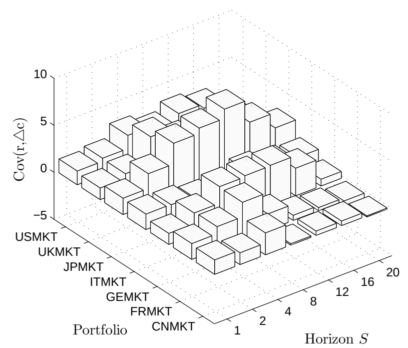
(c) Int. FX Portfolios, S=1 (excl. Japan)



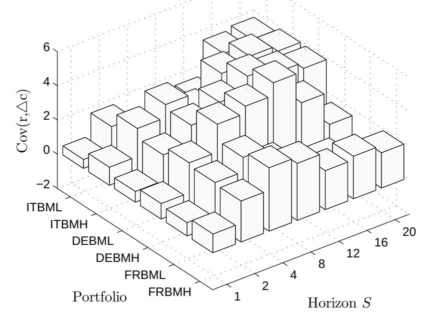
(d) Int. FX Portfolios, S=8

Notes: The Figure depicts pricing error plots for international bond and FX portfolios for horizons of long-run consumption growth of $S = 1$ and $S = 8$. The constant is included in the cross-sectional regression. Subfigures (a) and (b) use international bond portfolios as test assets, while subfigures (c) and (d) show pricing errors for international FX portfolios.

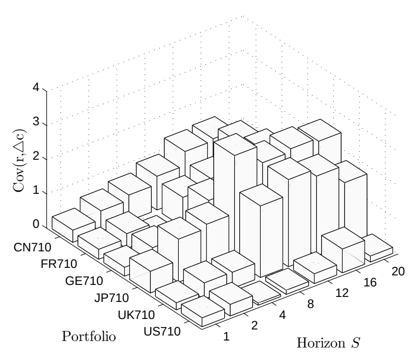
Figure 3: Covariances of Portfolios Returns with Long-run Consumption Growth



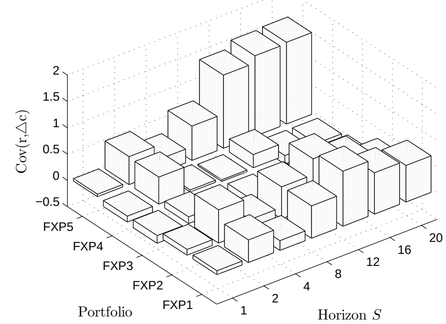
(a) Int. Equity Portfolios



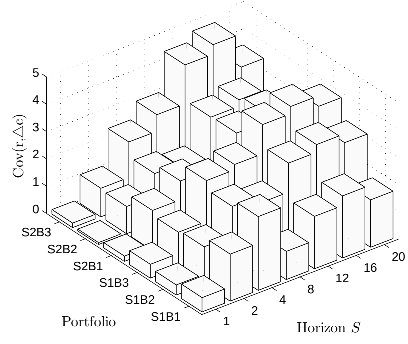
(b) Int. BM Portfolios



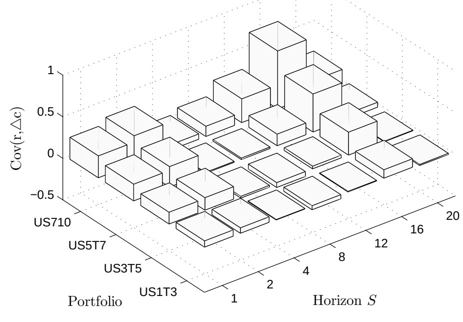
(c) Int. Bond Portfolios, (7-10 years)



(d) Int. FX Portfolios



(e) US, ME/BM Portfolios



(f) US, Bond Portfolios

Notes: The figure shows barplots of covariances of portfolio excess returns with discounted consumption growth for different asset returns and different consumption growth horizons $S = 1, 2, 4, 8, 12, 16, 20$. Subfigures (a) and (b) present plots for international aggregate stock market portfolios and international value/growth portfolios. Subfigures (c) and (d) present covariance plots for the entire set of international bond portfolios (with 7-10 years of maturity) and for 5 FX portfolios sorted by forward-discounts. Covariance plots for U.S. size and book-to-market sorted portfolios and bond portfolios (all maturities) are shown in subfigures (e) and (f).