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Abstract

We estimate long-run consumption-based asset pricing models using a comprehensive set of international test assets, including broad equity market portfolios, international value/growth portfolios, and international bond portfolios. We find that differences in returns across assets *within* a country are sometimes – and most prominently for the U.S. – better captured by the assets' exposure to long-run consumption risk as opposed to their exposure to one-period changes in consumption (the canonical consumption CAPM). *Across* countries, however, exposure to long-run consumption risk does not provide a better fit than the canonical consumption CAPM. Thus, when characterizing the cross-country distribution of returns, long-run consumption risk does not seem to play any particular role, even if long-run risk is important for explaining the cross section of expected returns in the U.S. Furthermore, we show that consumption growth is more predictable over short to medium-run horizons than over longer horizons and that empirical evidence of a declining risk aversion parameter estimate in long-run risk models has to be interpreted with care.

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 contemporaneous consumption risk. In its standard form, this model has failed on a grand scale (see, e.g., Breeden and Litzenberger, 1989 or Mankiw and Shapiro, 1986). 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 or used to explain other features of U.S. financial assets (see Daniel and Marshall, 1997; Parker, 2003; Bansal and Yaron, 2004; Parker and Julliard, 2003; Parker and Julliard, 2005; Hansen, Heaton, and Li, 2008; and Malloy, Moskowitz, and Vissing-Jørgensen, 2009).

In this paper, we use a consistent set of test assets from the G-7 countries during the 1973-2008 period to comprehensively evaluate whether the empirical success of the long-run risk model extends to an international asset pricing context. Our test assets include broad stock market portfolios, value/growth stock portfolios, and portfolios of bonds of different maturities. The broad set of test assets allows us to combine the assets in different ways: As cross sections of assets from one country only or as cross-country sets of assets.

First, we look at the countries one by one. We ask how well the models explain the joint distribution of returns on stock and bond portfolios *within* each of the countries. We find that exposures to long-run risk improve the performance of the model compared to the canonical consumption CAPM for some countries, in particular and most impressively for the U.S.[1] For instance, the canonical consumption CAPM that prices the joint cross section of U.S. stocks and bonds via their exposures to the one-period growth rate of U.S. per capital consumption generates a risk aversion coefficient as high as 134; the constant in the empirical moment function is significantly different from zero (implying that there is a constant mispricing of all assets); and the cross-sectional R^2 is 40%. Compare this to the

[1] Our empirical implementation of the long-run C-CAPM follows Malloy, Moskowitz, and Vissing-Jørgensen (2009). The Malloy, Moskowitz, and Vissing-Jørgensen (2009) procedure is based on 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, Moskowitz, and Vissing-Jørgensen (2009), we focus on the situation where the EIS is equal to one and the risk aversion parameter is freely estimated.

situation where we explain the cross-sectional distribution of returns on U.S. assets via their exposures to long-run consumption growth over, for instance, the next sixteen quarters. In that model, the estimated risk aversion coefficient is only half of what we find for the canonical C-CAPM, the constant in the empirical moment function is insignificant, and the R^2 is an impressive 85%.[2]

When estimating models for Canada, France, Germany, and the U.K., we find somewhat similar patterns and large cross-sectional R^2s, although the picture is not quite as impressive as for the U.S. sample when using data from Canada and Germany (for France and the U.K., though, the fit is remarkable; for the U.K., for instance, a R^2 of 80% and a risk aversion coefficient of 20 when using long-run risk). For Japan, allowance for long-run consumption exposure does not improve the fit of the model compared to the canonical C-CAPM. Based on the initial individual-countries estimations, we conclude that consumption-based models in many countries work astonishingly well on the bond and stock cross-sections investigated here, but also that there is considerable heterogeneity regarding the importance of long-run as opposed to short-run consumption risk.

We then proceed to the next natural step and try to jointly price the international cross-country distribution of returns. In this international asset pricing model, the level of average excess returns is explained by the assets' exposures to international risk factors. We examine two consumption risk factors: World-consumption growth and U.S. consumption growth.[3] Our results are consistent using either type of consumption. Indeed, our main finding is that the models are unable to price the cross-country distribution of returns successfully when using contemporaneous consumption growth, but also – and unlike in the U.S. data – when using exposures to long-run consumption growth (regardless of whether this is long-run U.S. or world-consumption growth). For instance, the cross-sectional R^2s remain low regardless of the horizon over which we measure the assets' consumption exposure, the estimates of the

[2]We stress that a cross-sectional R^2 of 85% for a model that prices both U.S. stocks and bonds with a single consumption-based factor is noteworthy in light of the recent findings of Koijen, Lustig, and van Nieuwerburgh (2009).

[3]Papers using U.S. consumption to test international consumption-based asset pricing models include Cumby (1990), Wheatley (1988), and Lustig and van Nieuwerburgh (2005), whereas papers that use world-consumption risk to price international assets include Sarkissian (2003) and Li and Zhong (2005). Obviously, the choice of consumption series depends upon the degree of integration of capital markets (see also Stulz, 1981). With perfect integration, all assets are priced according to their exposures to world risk factors, whereas local risk factors are relevant if capital markets are segmented. When examining U.S. and world consumption, we avoid assuming any particular degree of capital-market integration.

risk aversion parameter remain implausible, and the constant remains significant (which it should not). Hence, our overall conclusion in this paper is that the long-run model (and the canonical C-CAPM) does a poor job when used to explain the *international cross section* of returns, even when the distribution of returns *within* some countries (in particular the U.S.) is significantly better explained by the assets' exposures to long-run consumption risk.

In addition to this main result, we provide an interesting, more methodological, empirical result regarding the estimation of the risk aversion coefficient in long-run risk models. We find that the estimates of the risk aversion parameter seem to be much improved (i.e. are much lower) in long-run risk models when we do not allow for a constant in the empirical moment function:[4] While the estimate of the risk aversion coefficient is 355.83 in the standard canonical C-CAPM when there is no constant in the empirical moment function, it is "only" 46.54 when we expose the assets to consumption growth over the next twenty quarters. Moreover, the decline is almost monotonic in the horizon over which we measure the consumption exposure. This seems to indicate an improvement of the international long-run risk model over the canonical C-CAPM in the models that restrict the constant to zero. It turns out, however, that this seeming improvement stems mostly from inflating the variance of consumption growth with a growing horizon rather than an increase of the correlation of excess returns with consumption growth. To show this, we bootstrap consumption growth rates under an iid assumption and estimate the models using the original returns and the bootstrapped iid consumption growth rates. We find that in these simulated samples – where consumption growth data are not related to returns per construction – it is still relatively easy to obtain patterns closely resembling the ones in the original data. Hence, we conclude that the decline in the risk aversion parameter that we observe in the models where there is no constant should not be interpreted as a sign of success of the long-run risk model.

Our results are robust. First, they hold if using both U.S. and world consumption as the measure of consumption risk towards which the assets are assumed exposed, as mentioned. Second, we rely on the approximate linear stochastic discount factor of Malloy, Moskowitz, and Vissing-Jørgensen (2009) in our estimations. Parker and Julliard (2005) use a related, but exact and non-linear stochastic discount factor. Hence, we estimate models using the Parker

[4]More specifically, we estimate two kinds of models: One where following Malloy, Moskowitz, and Vissing-Jørgensen (2009), we allow for a constant in the empirical moment function and one where we a priori restrict the constant to zero, which is what it actually should be from a theoretical perspective. We find that the behavior of the risk aversion coefficient, at first sight, seems to depend upon the constant being zero or not.

and Julliard approach and find the same results. Third, it turns out that Japan is in some ways an "outlier". Hence, we estimate the asset pricing models using data for all countries except Japan. We report the same kind of results. We conclude that our overall finding that exposure to long-run consumption risk does not improve the performance of international consumption-based CAPMs over standard canonical consumption CAPMs seems robust.

The remaining part of the paper is structured as follows: In the next section, we describe the long-run risk model that we estimate. Section 3 presents the data. In section 4, we investigate the performance of the long-run risk C-CAPM when used to explain the joint cross section of U.S. stock and bond returns as well as similar tests for the within-country cross-sections of the remaining capital markets . Section 5 examines the performance of the model when used to explain cross-country differences in returns. Section 6 provides simulation-based evidence on the properties of the coefficient of relative risk aversion in the long-run risk model. In section 7, we present robustness checks on our basic findings. Section 8 concludes.

2. Asset pricing and long-run consumption risk

2.1. Theoretical framework

Our estimations are based on the empirical framework of Malloy, Moskowitz, and Vissing-Jørgensen (2009), henceforth MMVJ, which in turn builds on the theoretical work of Hansen, Heaton, and Li (2008), henceforth HHL. HHL and MMVJ assume that investors have recursive Epstein and Zin (1989) preferences:

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

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 an 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 is a vector of state variables, and G has eigenvalues less than one. Under these assumptions, HHL and MMVJ show that the log-linearized asset-pricing (Euler) equation for the return on an asset i over and above the risk-free rate is:

$$\mathrm{E}\left(r^i_{t+1} - r^f_{t+1}\right) + 0.5\sigma^2(r^i_{t+1}) - 0.5\sigma^2(r^f_{t+1}) \simeq (\gamma - 1)\,\mathrm{Cov}\left(r^i_{t+1} - r^f_{t+1}, \Sigma^{\infty}_{s=0}\beta^s \Delta c_{t+1+s}\right) \tag{3}$$

when the elasticity of intertemporal substitution in consumption is equal to one. In Eq. (3), $r^i_{t+1} = \ln(1 + R^i_{t+1})$ is the log-return on asset i, $r^f_{t+1} = \ln(1 + R^f_{t+1})$ is the log risk-free rate, and $0.5\sigma^2(\cdot)$ are the usual variance terms arising from the log-linearization. The key point to note 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. 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.[5]

Malloy, Moskowitz, and Vissing-Jørgensen (2009) and a related paper by Parker and Julliard (2005) focus on the performance of long-run risk models to explain the U.S. cross section of stock returns (twenty-five Fama-French portfolios). Both papers find that γ is generally estimated to be high when only contemporaneous consumption growth is 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

[5]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.

to be low and/or more precisely estimated.[6] 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. Thus, given the promising results provided in the extant literature estimating long-run risk models on the U.S. equity cross section of returns, it is natural to ask how well the model is able to explain international asset prices. Assessing the performance of the long-run risk model in an international asset pricing context is the main purpose of the current paper.

2.2. Empirical implementation

Our empirical implementation of the model by the generalized method of moments (GMM) builds on the moment conditions implied by Eq. (3). The empirical moment function with $f_{t+1} = \Sigma_{s=0}^{S-1} \beta^s \Delta c_{t+1+s}$, i.e. after having truncated the infinite sum of future consumption growth rates in Eq. (3) at horizon S, then reads:

$$h(\gamma, \alpha, \mu_{c,S}; \Theta_{t+1}) = \begin{bmatrix} \mathbf{r}^e_{t+1} + 0.5\sigma^2 - 0.5\sigma^2_f - \alpha\iota_N - (\gamma - 1)\, \mathbf{r}^e_{t+1} (f_{t+1} - \mu_{c,S}) \\ f_{t+1} - \mu_{c,S} \end{bmatrix} \tag{4}$$

where $\mathbf{r}^e_{t+1}$ is the vector of excess returns from the N test assets and σ^2 and σ^2_f denote vectors collecting the variances of the excess returns and the risk-free rate, respectively. In the empirical implementation, we use the unconditional return variance (computed for the full sample) as our estimate of σ^2.

As in Parker and Julliard (2005) and Malloy, Moskowitz, and Vissing-Jørgensen (2009), we estimate models where we include a constant (α) in Eq. (4), even if it should not be there from a theoretical perspective; in Eq. (3) there is no constant. 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.

[6] 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.

We estimate Eq. (4) using GMM. The GMM procedure finds the values of the parameters γ, α, and $\mu_{c,S}$ that best satisfy 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 pre-specified weighting matrix:

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

as in Parker and Julliard (2005) and Malloy, Moskowitz, and Vissing-Jørgensen (2009), such that the portfolios are given equal weight in the minimization.[7] 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. Furthermore, we present two measures of fit for each of our estimations: The cross-sectional R^2 and the Hansen and Jagannathan (1997) distance along with p-values from tests of whether the HJ-distances are statistically distinguishable from zero.[8]

3. Data

In this section, we briefly discuss the data we use. Summary statistics are provided in Table 1.

Consumption data. In order to investigate the long-run consumption CAPM in an international asset pricing context, a natural choice of consumption would be world consumption growth. Due to data availability, we limit ourselves to quarterly, private total consumption time series in the G-7 countries. The data are 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.[9] For robustness, we also estimate models using U.S. real per-capita consumption of non-durables and services, which is the series commonly used in the national asset pricing literature. Finally, we use the countries' own consumption growth series when we estimate the (long-run) C-CAPM for each country in isolation.

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

[8] For further details on the computation of the HJ-distance and the statistical tests of whether it is equal to zero, see Jagannathan and Wang (1996) and Parker and Julliard (2005).

[9] This is a common procedure that has also been used in other papers exploring international aspects of consumption-based asset pricing models, such as Sarkissian (2003), and Li and Zhong (2005), among others.

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. Harvey, 1991, Dumas and Solnik, 1995 and De Santis and 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 1973Q2-2007Q4.

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 (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 six 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.[10] For each non-U.S. market there is both a high book-to-market (value) and a low book-to-market (growth) portfolio available. Thus, in total we use sixteen portfolios (expressed in U.S. dollars) which cover the sample period 1975Q1-2007Q4.

We also include international bond returns. Our data are total returns on Merrill Lynch Government Bond Indices for the G-7 (excluding Italy) which are taken from Datastream.[11] For each country, we use four maturity categories: 1-3, 3-5, 5-7, and 7-10 years. Hence, in total, there are twenty-four portfolios at the quarterly frequency, and the sample period is 1986Q2-2007Q4.

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.

Summary statistics. Table 1 contains descriptive statistics for our test assets' excess returns. The table contains means, standard deviations, and Sharpe ratios of the test assets' excess returns, each of which are expressed in annualized percentage points. Furthermore, the statistics reported in the table include skewness, excess kurtosis, the autocorrelation co-

[10] Canada is omitted since the time series does not cover the full sample period in the Fama-French dataset.

[11] Italy is omitted since the time series does not cover the entire time period. Data for Italian bonds start in the 1990s.

efficient of first order as well as the co-skewness measure introduced by Harvey and Siddique (2000).[12]

– Insert Table 1 about here –

As shown in Panel A, annualized average excess returns in international stock markets range from about 5.5% (Japan) to about 10.5% (France), while Sharpe ratios range from about 0.23 (Japan) to about 0.44 (U.S.). Furthermore, Panel B shows that there is a clear tendency for value stocks to earn higher returns than growth stocks in international equity markets; Sharpe ratios of value stocks also tend to be higher than those of growth stocks. Panel C reports summary statistics for the excess returns on the 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] The means rise even faster with the maturity, leaving the Sharpe ratios higher for the portfolios of bonds with longer maturities.

Informational content of returns for future consumption. The major reason why Parker and Julliard (2005) and Malloy, Moskowitz, and Vissing-Jørgensen (2009) 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 tends to increase with S and it increases in the right way for the right assets.[14] To obtain an initial impression of our dataset, Table 2 presents results from simple regressions of the form:

$$\Sigma_{s=0}^{S-1}\beta^s\Delta c_{t+1+s} = \alpha + b_1 PC1_t + b_2 PC2_t + \varepsilon_{t+1+s} \tag{5}$$

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

$$\Sigma_{s=0}^{S-1}\beta^s\Delta c_{t+2+s} = \alpha + b_1 PC1_t + b_2 PC2_t + \varepsilon_{t+2+s} \tag{6}$$

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

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

[14] For instance, Parker and 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."

on the right-hand-side in columns I.B, II.B, and III.B, where Δc_{t+i} is the one-step change in world consumption from $t+i-1$ to $t+i$.[15] 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. (6). The explanatory variables we use are the first two principal components from the set of excess returns.

Hence, Eq. (6) can essentially be regarded as a forecasting regression for future long-run consumption growth. Cochrane (2007), page 284, explicitly asks for regressions as in Eq. (6), such that the predictability of future growth rates can be evaluated. And for good reason. The well-established dismal performance of the canonical C-CAPM stems from the fact that the correlation between consumption growth and returns is "too low" and that consumption growth is "too smooth" (Cochrane, 2005). Using long-run consumption growth naturally alleviates the latter point but not necessarily the former.[16] Thus, these regressions can serve as an initial check to investigate how considering long-run consumption risk can potentially improve upon the standard CCAPM.

– Insert Table 2 about here –

The left-hand side of Table 2 shows that the relation between asset returns and consumption growth (as given by the R^2) tends to be higher for horizons beyond $S = 1$ (the case of contemporaneous consumption growth). However, most importantly, Table 2 (right-hand side) shows that long-run consumption growth is, to some extent, predictable by the asset returns. This predictable component is quite sizeable, with predictive R^2s of 7-8% for stock markets. The existence of a predictable component is at odds with consumption growth being iid, but it constitutes a precondition for the potential success of an asset pricing model featuring long-run consumption risk.

However, the table also shows that there seems to be more predictability of consumption growth over relatively short horizons of one to four quarters, whereas the evidence for predictability is considerably weaker for longer horizons of eight to twelve quarters. This observation already suggests that "short- to medium-run" consumption growth may be a more successful factor in international asset pricing than long-run consumption growth.

[15] We only present the results for world consumption. Results for U.S. consumption are qualitatively similar.

[16] We elaborate on this point in more detail in section 6 below.

4. Results for within-country cross-sections

This section shows results that document that long-run consumption risk can help capture the cross-sectional distribution of returns *within* some countries, in particular the U.S. First, we discuss the results obtained when analyzing U.S. returns. Afterwards, we discuss the results obtained when analyzing the returns from the other countries.

4.1. Positive evidence: Results for the U.S.

Table 3 presents the results from the first asset pricing tests that we conducted. The table shows how exposures to U.S. consumption (in the upper part of the table) and world consumption (lower part of the table) capture the cross-sectional variation of average excess returns for different consumption-growth horizons, i.e. for different values of S. There are eleven assets to be priced: The U.S. market portfolio, the six size-/book-to-market portfolios, and the four bond portfolios. For each value of S, the table shows the estimate of the constant in the empirical moment function ($\widehat{\alpha}$) and the estimate of the risk aversion parameter ($\widehat{\gamma}$) with t-statistics based on Newey-West standard errors below coefficient estimates. We also show two measures of fit: The cross-sectional R^2 and the HJ-distance (and the associated probability values from tests of whether the HJ-distance is equal to zero).

– Insert Table 3 about here –

Looking at the results from the estimations of the canonical C-CAPM (the model using one-period change in U.S. consumption, i.e. $S = 1$, to price the assets) shows that the risk aversion coefficient is estimated to be very high: 133.60. Such high estimates of γ are well-known from the equity premium puzzle literature based on U.S. data.[17] 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. In addition, the constant is positive and significant. We use excess returns, so the constant should be zero. The constant, however, is estimated to 0.009, i.e. the average historical level of excess returns is approximately 4%

[17] A recent example is Lettau and Ludvigson (2009), page 260, who report an estimate of the risk aversion parameter of 118.

higher than the corresponding level implied by the model.[18] The R^2 is estimated to be 40%. This perhaps seems high at first glance, but, as we will see next, compared to other models, it is actually not overly impressive.

Now consider what happens if the horizon over which consumption growth is measured is increased, i.e. S is increased. Four things happen: The estimate of the risk aversion coefficient is lower (at $S = 16$, for instance, it is estimated at 65.91), the estimate is more precise (at $S = 16$, for instance, the t-statistic is equal to 1.82), the constant is insignificantly different from zero at $S = 16$, and, probably most noteworthy, the R^2 is a very high 85%. Compared to recent findings in the literature (Koijen, Lustig, and van Nieuwerburgh, 2009), a cross-sectional R^2 of 85% is high in a model that prices both U.S. stocks and bonds at the same time. All in all, we conclude that the cross-sectional distribution of excess returns on U.S. bonds and stocks is clearly better captured by the assets' exposures to consumption over longer periods and the implied economic parameters are more reasonable.[19]

We will later use world consumption to price the cross-country distribution of assets from many countries. To verify that the use of world-consumption data does not give rise to results that qualitatively differ much from those using local (in this case U.S.) consumption, the lower half of Table 3 also shows the results we get if using world consumption to price U.S. assets. The table reveals the same patterns as when using U.S. consumption: Better estimates of the risk aversion parameter when using changes in consumption over longer horizons (when using changes over short horizons, the risk aversion parameter is even estimated to be negative), the constant being significant at shorter horizons but insignificant at longer, higher cross-sectional R^2s at longer horizons, etc. Hence, the qualitative conclusions we draw on the basis of world-consumption risk are essentially the same as those we get using local U.S. consumption.

[18] The finding that consumption-based models are not consistent with the high level of excess returns is a common finding in the literature. On a related matter, estimates of the risk-free rate are often too high in this kind of model (see, for instance, Lettau and Ludvigson, 2001).

[19] Of course, a risk aversion coefficient of 65.91 is still high. The point is, though, that allowance for long-run consumption exposure reduces the estimate by approx. 50% compared to the value implied by the canonical C-CAPM.

4.2. Mixed evidence: Results for other individual countries

Table 4 shows the results we get if running the asset pricing tests for the other countries, treating each country on its own, i.e. estimating one asset pricing model for each country (for each value of S). For each country, there are seven assets: The market portfolio, two book-to-market portfolios, and four bond portfolios. We use the model to price the cross section of returns on the assets via the assets' exposure to local consumption (left columns in Table 4) and world consumption (right columns).

– Insert Table 4 about here –

The picture is mixed. There are some countries where long-run risk helps in pricing the cross section of assets. For instance, in the models estimated on data from, respectively, Canada, France, and U.K., the estimate of the risk aversion parameter is more sensible using exposure to long-run risk and the R^2s are higher. For Germany, the risk aversion coefficient is more reasonable using exposure to long-run consumption, but the R^2 is lower. Regarding the constant: If it is significant (insignificant), it generally remains so regardless of the value of S for Canada, France, Germany, and the U.K. In other words, allowing for long-run exposures does not change the significance of the constant. For Japan, exposure to long-run risk clearly describes the cross-sectional distribution of returns worse than exposure to short-run consumption growth does.

All in all, we conclude from this and the previous section that the characterization of the cross-sectional variation of U.S. asset returns is much improved if explaining this distribution via the assets' exposure to long-run risk. Hence, for the U.S., long-run consumption risk is indeed relevant. For other countries, the story is not so clear: There are some countries where long-run risk helps, though to a lesser extent than in the U.S. data, while there are other countries where this is not the case. For Japan, the conclusion is that long-run risk even worsens the performance of the consumption-based model.

5. Negative evidence: Results for cross-country cross-sections

If an asset pricing model really works, it should, of course, price the cross-sectional distribution of returns on all assets and not only specific subsets of returns. With this motivation, we evaluate now whether the long-run risk model is better able to capture the cross-country distribution of returns compared to the standard consumption-based asset pricing models.

We use the series for G-7 consumption growth, calculated by the IMF, as our measure of the consumption risk factor. In order to make sure that our results are not driven by our choice of consumption series, we also estimate the models using U.S. consumption instead of G-7 consumption and present these results as robustness checks in Section 7. At this point, we can already mention, though, that we find the same qualitative results in these robustness tests as in our baseline case. In addition, Tables 3 and 4 show that the differences between the results we get when using world consumption instead of local consumption were not qualitatively large.

We present the main results of this paper in Table 5. The outline of the table follows the outline from Tables 3 and 4, i.e. we present the estimates of the constant, the risk aversion parameter, the R^2, and the HJ-distance, as well as relevant statistical measures (t-statistics and probability values). We do this for the different asset classes, as well as for the full model using all data. Taking the full model that tries to price all assets first, Panel D in Table Table 5 shows that (i) the constant is significant for all horizons, (ii) the R^2s increase a little with the horizon, but there is no impressive improvement in the cross-sectional fit, (iii) the HJ-distance measure is, if anything, even higher at the longer horizons, and (iv) even if there is some tendency for the risk aversion coefficients to be estimated to be lower when exposing the returns to longer-horizon consumption risks, the risk aversion coefficients are very imprecisely estimated. Hence, the table shows, in essence, that exposures to consumption risk cannot explain the variation across countries in excess returns, whether the exposure is measured using one-period changes in consumption growth rates (the canonical C-CAPM) or using long-horizon consumption risk. This result is very different from the result in the literature (based on U.S. data) that long-run risk helps: We find that it does so in the U.S. data (as shown above), but not in international data.

– Insert Table 5 about here –

Panels A, B, and C help us understand whether consumption risk can explain the returns within some international asset classes. Let's take the most positive results first. For the bond portfolios, there are some indications that exposure to long-run risk helps. The risk aversion parameter is negative when using exposure to short-horizon consumption changes, but positive (even if insignificant) when using exposure to long-run risk. Likewise, the R^2 is higher at the longer horizons. The constant mispricing, however, is significant both at short and long horizons. Long-run consumption exposure, thus, helps somewhat when explaining the cross-country distribution of returns from bonds. However, the cross-country distribution of returns from stocks is not at all well-captured, whether using exposure to short- or long-run consumption growth. Most clearly, the risk aversion parameters are estimated to be negative at all horizons (even if insignificant). Moreover, the constants are generally significant (or close to being so) at the longer horizon, and even if the cross-sectional R^2s are somewhat higher for the longer horizons, they are low compared to the values presented in Tables 3 and 4.

What drives this result? Figure 1 depicts the covariances of (selected) test asset returns with consumption growth measured over different horizons. This figure helps improve understanding as to why the long-run consumption-based approach fails in rationalizing return differences across countries. For the case of international equity indices, subfigure (a) shows that exposure to consumption growth often tends to rise when moving from $S = 1$ to horizons of about $S = 8, 12$ and then often declines afterwards. Most importantly, however, the magnitudes of the covariances *across* countries illustrates why there is no support for the long-run risk approach in explaining the cross-country variation in average returns. Recall from the discussion in section 3 and Table 1 that the French equity market had the highest return over our sample period while the Japanese market had performed the worst. As Figure 1 shows, however, Japanese stock returns have actually had the highest exposure to long-run consumption risk, whereas the high return on the French stock market does not seem to reflect compensation for a particularly high exposure to long-run consumption risk, which it should if the long-run consumption risk story is to describe returns. The trouble with the long-run risk approach in accounting for cross-country differences in returns also holds for international bond returns (subfigure c). Subfigure b, however, shows some signs of success of the model for explaining return differences between value and growth portfolios within individual countries (consistent with our discussion in section 4). All in all, this diagnostic exercise provides intuition about the dimensions along which the long-run risk model works and fails.

– Insert Figure 1 about here –

Figure 1 also helps us understand why we often, and in particular when S is low, obtain unreasonable estimates of γ in Table 5 when a constant is included in the empirical moment function. To explain this, bear in mind that there is not much cross-sectional variation in the covariances with long-run consumption growth in the case of the international cross-sections (Figure 1 shows that the covariances are basically the same across the different assets for low values of S). Intuitively, when expressed in the language of a traditional cross-sectional regression framework, this lack of cross-sectional variation gives rise to multicollinearity type problems as the risk factor loadings act as a "second constant" in the cross-sectional regression when a constant is included. Hence, the risk aversion parameter is not identified in a meaningful way in this case. A similar problem occurs, for instance, when estimating a cross-sectional regression with a constant for a traditional CAPM on the twenty-five Fama-French portfolios since market betas do not show much variation across the 25 Fama-French portfolios (see, e.g., Lettau and Ludvigson, 2001).

6. Evidence that seems positive, but is not

Following Malloy, Moskowitz, and Vissing-Jørgensen, 2009, we allow for a constant mispricing of all assets by including a constant in the empirical moment function. Theoretically, the constant should not be there, of course. Hence, we also estimate the models without the constant. Table 6 contains the results.

– Insert Table 6 about here –

At first sight, the results seem to provide considerably more reasonable estimates of the risk aversion parameter when using exposure to long-run consumption risk. Looking at the results obtained when we use all assets, we see that the risk aversion parameter is estimated at 355.83 when using exposures to quarterly consumption growth changes but "only" 53.38 when using exposures to changes in consumption over the next twenty quarters. The decline in the values of the estimated parameters is also clearly seen in the other sub-portfolios (the broad equity market indices, the value/growth portfolios, and the bond portfolios). This finding seems to indicate that there may be (at least some) success for the long-run risk theory.

However, the fact that the R^2s do not increase and the HJ-distances, if anything, increase with the horizon makes us sceptical about whether there really is an underlying economic reason for the pattern of declining risk aversion parameter estimates provided in Table 6.[20] Indeed, as we argue below, it seems reasonable to assume that this finding is mainly driven by a mechanical effect when using sums of future consumption growth rates as a risk factor in these empirical exercises.

To state this more explicitly, consider the following expression for the risk aversion parameter obtained from rearranging Eq. (3) and abstracting from the $0.5\sigma(\cdot)$ terms:

$$\gamma \simeq 1 + \frac{E(r_{t+1}^{e,i})}{\rho(r_{t+1}^{e,i}, f_{t+1}^{S})\sigma(r_{t+1}^{e,i})\sigma(f_{t+1}^{S})} \tag{7}$$

where $r_{t+1}^{e,i}$ is the excess return of asset i, f_{t+1}^{S} denotes demeaned long-run consumption over horizon S and $\rho(\cdot,\cdot)$ is the correlation coefficient. In this setting, two things can bring down the risk aversion coefficient, γ, when the horizon, S, increases: (a) an increase in the correlation of excess returns with consumption growth and/or (b) a more volatile consumption risk factor. In economic terms, one would hope that long-run consumption risk increases the correlation part. This is the case, as shown in section 4, for the U.S. joint cross section of bond and stock returns, and this is also what Parker and Julliard (2005) and Malloy, Moskowitz, and Vissing-Jørgensen (2009) have demonstrated for the twenty-five Fama-French portfolios.

For the international cross section of assets that we study in this paper, however, correlations between returns and consumption growth rates are actually quite flat – on average across assets – and after an increase for short- and medium-term horizons, they tend to decrease over longer horizons (see Table 2). Hence, in this case, the decline of γ with increasing horizon

[20] Of course, one should take care not put too much emphasis on the R^2s as there is no constant in the empirical moment function.

S can mostly be attributed to an increase in the volatility of the risk factor, which arises naturally when summing up one-step consumption growth rates (which are approximately iid) over longer horizons.

In order to investigate this issue more closely, we proceeded as follows: Assuming that consumption growth rates are iid, we generated 1,000 samples of artificial consumption growth series from the original one-period consumption growth series. We then formed cumulative consumption growth rates from these artificial consumption growth series and ran our estimations (using the original excess returns) for each of the 1,000 artificial samples. Given that the consumption growth rates are iid, and thus unrelated to returns per construction, we should not expect to see any systematic differences between the results using short- and long-horizon consumption exposure. We do see systematic differences, though. Figure 2 and Table 7 contain the results. Consider Figure 2 first.[21] The figure reveals a pattern almost identical to the one shown in Table 6, i.e. declining estimates of the risk aversion parameter when using exposures to long-run risk for all three subportfolios. However, we know that the consumption growth series used to generate the pattern in Figure 2 is unrelated to returns, and hence there is no extra economic content in the consumption growth series of the long-run consumption growth series in the bootstrapped data.

– Insert Figure 2 about here –

To provide a more detailed picture, Table 7 lists the fraction of times that the risk aversion parameter is, respectively, 25%, 50%, and 75% smaller using twenty quarters of consumption growth rates compared to the standard case of quarterly consumption growth changes for three different cases: When $\gamma_{S=1}$ is positive, when $\gamma_{S=1}$ and $\gamma_{S=20}$ are both positive, and when all $\gamma_{S=1}$, $\gamma_{S=2}$,...., $\gamma_{S=20}$ are positive. Consider, for example, case 3 in Table 7, where the requirement is that all estimates of the risk aversion coefficient should be positive. In this case, we see that in 70% of the estimations, the value of $\gamma_{S=20}$ is at least 75% smaller than the value of $\gamma_{S=1}$ (using the portfolios of international market indices). Likewise, this was the case in 57.3% of the estimations using the simulated data for the international bond indices and in 33.3% of the estimations using the data for the value/growth portfolios.

[21]For each of the 1,000 artificial samples, the LR-CCAPM is estimated for horizons $S = 1, ..., 20$ (no constant in the empirical moment function). The mean risk aversion coefficient (for the 1,000 samples) for which $\gamma_{S=1,...,20}$ are jointly positive is plotted for different horizons.

All in all, we conclude from this section that even though the estimates of the risk aversion parameters get smaller when using exposure to long-horizon risk in models where we restrict the constant in the empirical moment function to being zero (as it theoretically should be), one should be very careful in interpreting this phenomenon as a "successful" aspect of the model. Indeed, in simulated data where long-run consumption growth rates do not contain more information than short-run consumption growth rates by construction, we quite often find the same pattern.

– Insert Table 7 about here –

7. Robustness

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

U.S. consumption data. In order to evaluate whether our results from the international cross-country asset pricing tests are sensitive to our choice of consumption, this section presents estimates from models using U.S. consumption (of nondurables and services) and compares them with the results we get when we use the world consumption series. We present both results that include and exclude a constant in the empirical moment function. The results appear in Table 8.

– Insert Table 8 about here –

The main conclusion is that we find basically the same results when using U.S. consumption growth to measure consumption risk as the ones reported earlier. First, consider the results we find if including a constant in the empirical moment function. We find that there is generally not an improvement in the estimates of the risk aversion coefficient (with the possible exception of the bond portfolios), that the constant is often significant, and that the cross-sectional R^2s increase for some portfolios (international stock market indices) but not for others (value/growth and the bond portfolios). Next, consider the results we get if the constant is excluded in the moment function. Like the results reported in Table 6 that were based on world consumption as the risk factor, the results in Table 8 reveal a declining value of the estimate of the risk aversion parameter if we exclude the constant. However, as Figure 2 and Table 7 show, we find the same pattern relatively often in artificially generated samples of bootstrapped iid consumption growth rates. Overall, the results provided in this paper are not sensitive to our choice of consumption series; qualitatively, we find the same results whether we use world consumption or U.S. consumption to measure consumption risk in international portfolios.

Comparison with Parker and Julliard. The asset pricing equation of Eq. (3) is approximate, as it is a log-linearized one. Parker and 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 and 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 return on any asset, i, between periods t and $t+1$, which results in:

$$E_t\left[m^S_{t+S}\left(R^i_{t+1} - R^f_{t,t+1}\right)\right] = 0, \tag{8}$$

where $m^S_{t+S} = (C_{t+S}/C_t)^{-\gamma} R^f_{t+1,t+S}$ after assuming that $U(C_t)$ is the standard power utility function $U(C_t) = C_t^{1-\gamma}/(1-\gamma)$. Eq. (8) 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.[22]

Table 9 shows the results of these tests.[23] Qualitatively, we find the same results as those reported above when we use the MMVJ approach: When there is a constant in the empirical moment function, there is no clear pattern in the estimates of the risk aversion coefficients,

[22]PJ test their model on the twenty-five Fama-French portfolios. Like MMVJ, Parker and 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 (2003) uses similar approaches to show that γ is estimated to be lower when S is larger.

[23]Our results in Table 9 are based on the linearized one-factor version of the PJ model.

but the R^2s sometimes increase with S. On the other hand, when there is no constant in the empirical moment function, the estimates of the risk aversion parameter decline in value the higher S is, but the R^2s do not increase. One small notable difference to the previous results, though, is that the estimates of the risk aversion parameter are more precise in Table 9 when there is no constant in the empirical moment function (the γs are all significantly different from zero in the no-constant case), whereas this is less clear in Table 6.

– Insert Table 9 about here –

Excluding Japan. As mentioned in section 3, the Japanese stock market has performed the worst in our sample of stock markets. In addition, Table 4 shows that Japan was the only country for which the cross-sectional fit (measured via the R^2) was much lower at longer horizons. Thus, one may ask whether the results we report for the international portfolios are due to the Japanese stock market behaving in a "strange way". Table 10, which shows results from estimations where we have excluded Japan, verifies that this is not the case, as the table shows that we do not find more convincing results for the consumption-based asset pricing model when using exposure to long-run consumption growth to price the assets, even if excluding assets from Japan.

– Insert Table 10 about here –

8. Conclusion

Recent results in the literature show that consumption-based asset pricing models that explain the differences in returns across U.S. assets via their exposures to U.S. long-run consumption growth bring down estimates of the risk aversion parameter to more reasonable levels and increase the cross sectional fit. We have investigated whether a similar success of the long-run risk model can be found in an international asset pricing context. We find the model to be successful for some countries when analyzing each country individually.[24] Even more inter-

[24] In particular, we are impressed by the explanatory power of long-run consumption risk for the cross section of bonds and stocks in the U.S. given that Koijen, Lustig, and van Nieuwerburgh (2009) argue that it is difficult to find models that can price the cross section of both U.S. stocks and bonds.

esting, though, we also find that consumption-based CAPMs cannot explain the international cross-country distribution of returns, whether we use the assets' exposures to single-period changes in U.S. or world consumption (the canonical consumption CAPM) or exposures to multi-period changes in U.S. or world consumption (the long-run risk models). Last, we show that the improvement in the estimate of the risk aversion coefficient in models where there is no constant in the empirical moment function is not surprising, as it follows more or less mechanically in a setting where consumption growth has no sensible cross-sectional correlation with returns, and is rather easy to replicate in a simulation where consumption growth does not matter at all for returns by construction.

Our findings raise the question of how the models can be improved in order for them to perform just as well as the models that work on data from the U.S. Possible extensions that could help reach this goal would be estimations of models that allow for non-perfect consumption risk sharing, as our findings also imply that even when financial markets are highly integrated today (Bekaert, Harvey, Lundblad, and Siegel, 2008), some countries' idiosyncratic consumption risk is not shared perfectly between the countries:[25] If all consumption risk was perfectly shared in the world, a single world consumption stochastic discount factor should price the cross-country distribution of returns (Stulz, 1981). As we have shown, though, this is not the case even when we have chosen a model that works very well in some countries. Hence, one potential avenue for future research is to allow for imperfect consumption risk sharing, like in Sarkissian (2003). One could also consider allowing for time variation in the volatility of consumption growth, as in the original paper of Bansal and Yaron (2004). Both of these extensions would make the model more flexible and thus more likely to fit returns better. We leave these exciting extensions to future research.

[25] Prasad, Kose, and Terrones (2003, 2009) show that even though consumption risk sharing has increased in developed countries since the late 1990s, it is still far from perfect.

References

Bansal, R., A. Yaron, 2004. Risks for the Long Run: A Potential Resolution of Asset Pricing Puzzles. Journal of Finance 59(4), 1481–1509.

Bekaert, G., C. Harvey, C. Lundblad, S. Siegel, 2008. What Segments Asset Markets?. Working Paper, Columbia University.

Breeden, D. T., R. Litzenberger, 1989. Empirical Tests of the Consumption-Oriented CAPM. Journal of Finance 44.

Cochrane, J. H., 2005. Asset Pricing. Princeton University Press, Princeton, NJ.

——— , 2007. Financial Markets and the Real Economy. in *Handbook of the Equity Premium*, ed. by R. Mehra. Elsevier pp. 237–325.

Cumby, R., 1990. Consumption Risk and International Asset Returns. Journal of International Money and Finance 9(2), 182–192.

Daniel, K., D. Marshall, 1997. Equity-premium and Risk-free Rate Puzzles at Long Horizons. Macroeconomic Dynamics 1, 452–484.

De Santis, G., B. Gerard, 1997. International Asset Pricing and Portfolio Diversification with Time-Varying Risk. Journal of Finance 52(5), 1881–1912.

Driessen, J., B. Melenberg, T. Nijman, 2003. Common Factors in International Bond Returns. Journal of International Money and Finance 22(5), 629–656.

Dumas, B., B. Solnik, 1995. The World Price of Foreign Exchange Risk. Journal of Finance 50(2), 445 – 479.

Epstein, L., S. Zin, 1989. Substitution, Risk Aversion, and the Temporal Behavior of Consumption Growth and Asset Returns I: A Theoretical Framework. Econometrica 57, 937–969.

Fama, E. F., K. R. French, 1993. Common Risk Factors in the Returns on Stocks and Bonds. Journal of Financial Economics 33, 3–56.

——— , 1998. Value Versus Growth: The International Evidence. Journal of Finance 53(6), 1975 – 1999.

Hansen, L. P., J. C. Heaton, N. Li, 2008. Consumption Strikes Back?: Measuring Long-Run Risk. Journal of Political Economy 116(2), 260–302.

Hansen, L. P., R. Jagannathan, 1997. Assessing Specification Errors in Stochastic Discount Factor Models. Journal of Finance 52(2), 557–590.

Harvey, C. R., 1991. The Term Structure and World Economic Growth. Journal of Fixed Income 1, 4–17.

Harvey, C. R., A. Siddique, 2000. Conditional Skewness in Asset Pricing Tests. Journal of Finance 55, 1263–1295.

Jagannathan, R., Z. Wang, 1996. The Conditional CAPM and the Cross-Section of Expected Returns. Journal of Finance 51(1), 3–53.

Koijen, R. S., H. Lustig, S. van Nieuwerburgh, 2009. The Cross-Section and Time-Series of Stock and Bond Returns. Working Paper, Chicago Booth.

Lettau, M., S. Ludvigson, 2001. Ressurrecting the (C)CAPM: A Cross-Sectional Test when Risk Premia Are Time-Varying. Journal of Political Economy 109, 1238–1287.

——— , 2009. Euler Equation Errors. Review of Economic Dynamics 12(6), 255–283.

Li, Y., M. Zhong, 2005. Consumption Habit and International Stock Returns. Journal of Banking and Finance 29(3), 579 – 601.

Lustig, H. N., S. G. van Nieuwerburgh, 2005. Housing Collateral, Consumption Insurance, and Risk Premia: an Empirical Perspective. Journal of Finance 60(3), 1167–1219.

Malloy, C. J., T. J. Moskowitz, A. Vissing-Jørgensen, 2009. Long-Run Stockholder Consumption Risk and Asset Returns. Journal of Finance (forthcoming).

Mankiw, G. N., M. D. Shapiro, 1986. Risk and Return: Consumption Beta versus Market Beta. Review of Economics and Statistics 68, 452–459.

Parker, J. A., 2003. Consumption Risk and Expected Stock Returns. American Economic Review 93(2), 376–383.

Parker, J. A., C. Julliard, 2003. Consumption Risk and Cross-Sectional Returns. NBER Working Paper 9538 National Bureau of Economic Research Cambridge, MA.

——— , 2005. Consumption Risk and the Cross Section of Expected Returns. Journal of Political Economy 113(1), 185 – 222.

Prasad, A., A. Kose, M. Terrones, 2003. How Does Globalization Affect the Synchronization of Business Cycles?. American Economic Review (Papers & Proceedings) pp. 57–62.

——— , 2009. Does Financial Globalization Promote Risk Sharing?. Journal of Development Economics (fothcoming).

Sarkissian, S., 2003. Incomplete Consumption Risk Sharing and Currency Risk Premiums. Review of Financial Studies 16(3), 983 – 1005.

Stulz, R. M., 1981. A Model of International Asset Pricing. Journal of Financial Economics 9(4), 383 – 406.

Wheatley, S. M., 1988. Some Tests of International Equity Integration. Journal of Financial Economics 21(2), 177 – 212.

Table 1: Descriptive Statistics of Portfolio Returns

Panel A: International Equity Portfolios (1973Q2-2007Q4)							
Portfolio	MEAN	STD	SR	SKW	KURT	COSK	AC1
Canada	7.102	18.301	0.388	0.065	-0.016	-0.078	0.020
France	10.516	24.262	0.433	0.357	1.142	0.056	0.181
Germany	7.679	20.431	0.376	0.297	0.394	-0.206	0.094
Italy	8.318	28.112	0.296	1.073	3.933	0.078	0.144
Japan	5.477	23.731	0.231	0.542	0.973	0.209	0.218
United Kingdom	9.666	22.937	0.421	0.495	1.865	-0.104	-0.063
United States	6.136	14.087	0.436	-0.004	0.838	-0.166	0.057
Panel B: International Value/Growth Portfolios (1975Q1-2007Q4)							
Portfolio	MEAN	STD	SR	SKW	KURT	COSK	AC1
France (high BM)	16.288	29.131	0.559	0.199	1.113	-0.126	-0.012
France (low BM)	9.459	24.163	0.391	0.204	0.954	-0.179	-0.001
Germany (high BM)	14.158	24.142	0.586	0.400	2.123	-0.278	0.018
Germany (low BM)	8.421	23.164	0.364	-0.250	0.932	-0.229	0.029
Italy (high BM)	9.125	32.803	0.278	1.236	4.386	-0.124	0.064
Italy (low BM)	9.059	29.001	0.312	0.997	1.959	-0.076	0.078
Japan (high BM)	13.821	26.358	0.524	0.097	0.034	0.000	0.058
Japan (low BM)	2.857	26.035	0.110	0.323	0.895	0.159	0.043
United Kingdom (high BM)	14.527	25.637	0.567	2.106	12.196	0.100	-0.088
United Kingdom (low BM)	11.106	24.071	0.461	2.075	11.861	0.112	-0.043
United States S1B1	9.343	26.896	0.347	0.042	0.365	0.108	-0.072
United States S1B2	15.041	20.579	0.731	-0.165	1.238	0.050	-0.091
United States S1B3	16.923	21.486	0.788	0.212	2.439	-0.021	-0.074
United States S2B1	7.774	18.119	0.429	-0.178	0.325	0.182	0.005
United States S2B2	9.915	15.364	0.645	-0.490	0.908	0.080	-0.043
United States S2B3	11.095	15.846	0.700	-0.158	1.792	0.005	-0.028
Panel C: International Bond Portfolios (1986Q1-2007Q4)							
Portfolio	MEAN	STD	SR	SKW	KURT	COSK	AC1
Canada (1-3)	4.627	7.066	0.655	0.405	0.931	0.181	-0.089
Canada (3-5)	5.659	8.200	0.690	0.325	0.447	0.188	-0.157
Canada (5-7)	6.166	9.766	0.631	0.440	0.442	0.348	-0.240
Canada (7-10)	6.635	10.028	0.662	0.228	0.440	0.212	-0.211
France (1-3)	4.626	10.427	0.444	-0.187	-0.504	0.500	0.018
France (3-5)	5.597	10.780	0.519	0.032	-0.359	0.569	0.016
France (5-7)	6.212	10.906	0.570	-0.017	-0.509	0.551	0.013
France (7-10)	6.746	11.522	0.586	0.221	-0.066	0.576	-0.004
Germany (1-3)	4.037	11.326	0.356	-0.111	-0.089	0.504	-0.022
Germany (3-5)	4.798	11.729	0.409	0.025	-0.107	0.527	-0.027
Germany (5-7)	5.363	12.057	0.445	0.122	-0.133	0.545	-0.046
Germany (7-10)	5.361	12.437	0.431	0.164	-0.187	0.551	-0.082
Japan (1-3)	1.798	13.549	0.133	0.908	1.138	0.585	0.052
Japan (3-5)	2.923	14.073	0.208	0.987	1.449	0.558	0.042
Japan (5-7)	3.794	14.928	0.254	1.003	1.586	0.537	0.017
Japan (7-10)	4.494	16.092	0.279	1.012	1.883	0.526	-0.042
United Kingdom (1-3)	5.383	10.863	0.496	0.493	1.698	0.791	-0.145
United Kingdom (3-5)	5.852	11.902	0.492	0.812	2.129	0.799	-0.179
United Kingdom (5-7)	6.443	12.968	0.497	0.863	2.338	0.802	-0.200
United Kingdom (7-10)	6.796	13.943	0.487	0.947	2.678	0.814	-0.212
United States (1-3)	1.616	2.020	0.800	0.188	0.192	0.288	0.141
United States (3-5)	2.656	4.024	0.660	0.094	0.102	0.285	0.051
United States (5-7)	3.253	5.180	0.628	0.108	0.147	0.303	0.037
United States (7-10)	3.628	6.407	0.566	0.041	0.254	0.304	0.034

Notes: This 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, the U.K. and the U.S.). Panel B includes descriptives for international value/growth portfolios (France, Germany, Italy, Japan, U.K.) and six U.S. size and book-to-market portfolios. Panel C reports results for the excess returns of international bond indices (Canada, France, Germany, Italy, Japan, the U.K. and the U.S.). Means (MEAN), standard deviations (STD), and Sharpe ratios (SR) are expressed in annual terms (in %). SKW denotes skewness and KURT excess kurtosis. Co-skewness (COSK) is computed according to Harvey and Siddique (2000) and AC1 denotes the autocorrelation correlation coefficient of first order.

Table 2: 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	1.985	2.472	4.544	2.929	3.148	PC1	0.469	1.390	1.936	0.196	0.888
	(3.63)	(2.45)	(3.23)	(1.09)	(0.87)		(0.67)	(1.38)	(1.29)	(0.06)	(0.24)
PC2	-0.864	-1.802	-4.684	-4.634	-3.644	PC2	-0.932	-2.098	-3.838	-3.582	-3.027
	(-1.86)	(-2.38)	(-3.04)	(-1.85)	(-1.45)		(-1.76)	(-2.26)	(-2.55)	(-1.55)	(-1.31)
R^2	9.97	8.95	17.86	5.48	2.79	R^2	2.41	6.18	7.93	2.43	1.23

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	0.553	2.307	3.496	3.615	4.137	PC1	1.739	2.326	3.436	2.994	4.045
	(1.07)	(2.62)	(3.07)	(2.08)	(1.76)		(2.39)	(2.45)	(3.14)	(1.57)	(1.54)
PC2	-0.417	-0.518	0.535	0.956	1.124	PC2	-0.095	0.370	0.885	0.433	1.246
	(-0.84)	(-0.81)	(0.40)	(0.39)	(0.32)		(-0.22)	(0.53)	(0.69)	(0.18)	(0.36)
R^2	1.19	5.80	5.46	2.47	2.08	R^2	7.07	5.60	5.53	1.65	2.05

III.A. International Bond Portfolios						**III.B. International Bond Portfolios**					
Horizon	1	2	4	8	12	Horizon	1	2	4	8	12
PC1	0.362	0.946	0.929	2.334	2.183	PC1	0.561	-0.101	1.044	2.163	1.810
	(0.91)	(1.33)	(1.03)	(1.42)	(1.01)		(1.11)	(-0.16)	(1.24)	(1.47)	(1.09)
PC2	-0.260	-0.890	-1.418	-2.132	-1.115	PC2	-0.649	-1.073	-1.358	-1.229	-0.839
	(-0.75)	(-1.60)	(-1.93)	(-1.64)	(-0.83)		(-1.60)	(-2.10)	(-1.89)	(-1.04)	(-0.58)
R^2	1.22	5.26	4.85	7.32	3.23	R^2	4.60	3.61	5.31	4.84	2.42

IV.A. All Portfolios						**IV.B. All Portfolios**					
Horizon	1	2	4	8	12	Horizon	1	2	4	8	12
PC1	-0.068	0.692	-0.006	-0.346	-0.342	PC1	0.757	0.361	-0.017	0.070	-0.473
	(-0.16)	(1.06)	(-0.00)	(-0.16)	(-0.13)		(1.59)	(0.40)	(-0.01)	(0.03)	(-0.18)
PC2	-0.175	-0.649	-0.600	-2.300	-3.103	PC2	-0.460	-0.262	-1.274	-2.835	-2.286
	(-0.38)	(-0.85)	(-0.53)	(-1.29)	(-1.33)		(-1.05)	(-0.45)	(-1.41)	(-1.80)	(-1.11)
R^2	0.16	2.44	0.39	2.11	2.38	R^2	4.25	0.54	1.79	3.28	1.39

Notes: This table provides regression results for 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 $\Sigma_{s=0}^{S-1} \beta^s \Delta c_{t+1+s}$ on the first two principal components of the portfolio returns R^e_{t+1}. Panel B presents results of the forecasting regression $\Sigma_{s=0}^{S-1} \beta^s \Delta c_{t+2+s}$ regressed on the first two principal components of the portfolio returns R^e_{t+1}. Standard errors are based on the adjustment by Newey and West (1987) with $S + 1$ lags.

Table 3: Within Country Cross-Sections: United States - Joint Bond and Value/Growth Cross Section with World and Local (Long-run) Consumption Growth

U.S. Cross-Section: Local Consumption							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.009	0.007	0.006	0.001	0.003	0.004	0.009
	(3.11)	(1.47)	(1.61)	(0.22)	(0.73)	(1.04)	(4.54)
$\hat{\gamma}$	133.60	171.53	85.04	112.64	75.09	65.91	37.55
	(0.95)	(0.71)	(1.09)	(1.07)	(1.64)	(1.82)	(1.68)
R^2	0.40	0.30	0.59	0.91	0.83	0.85	0.52
HJ-dist	0.65	0.65	0.63	0.53	0.51	0.60	0.63
	(0.00)	(0.00)	(0.01)	(0.20)	(0.14)	(0.01)	(0.00)
U.S. Cross-Section: World Consumption							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.019	0.020	0.009	0.014	0.010	0.008	0.009
	(2.32)	(2.72)	(1.61)	(2.60)	(1.87)	(1.37)	(2.47)
$\hat{\gamma}$	-224.48	-97.46	135.36	74.25	69.50	75.87	49.19
	(-1.17)	(-0.94)	(0.82)	(0.82)	(0.98)	(1.44)	(0.87)
R^2	0.22	0.05	0.32	0.12	0.32	0.46	0.45
HJ-dist	0.65	0.66	0.66	0.66	0.66	0.67	0.66
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Notes: This table reports estimation results for the LR-CCAPM for the within-country cross section of the United States (joint pricing of six value/growth portfolios and five government bond portfolios). Estimation results are presented using discounted long-run world and local 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 pre-specified weighting matrix. The reported estimation results are for the intercept α, the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey and West (1987) with $S + 1$ lags in parentheses), the cross-sectional R^2, and the HJ distance (simulation-based p-values in parentheses).

Table 4: Within Country Cross-Sections: International Markets - Joint Bond and Value/Growth Cross Section with World and Local (Long-run) Consumption Growth

Canada: World Consumption					**Canada: Local Consumption**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.020	0.009	0.013	0.012	$\hat{\alpha}$	0.020	0.010	0.013	0.012
	(2.42)	(1.05)	(2.16)	(2.31)		(2.59)	(1.29)	(2.44)	(2.75)
$\hat{\gamma}$	-96.51	76.58	25.88	19.09	$\hat{\gamma}$	-42.78	33.54	12.88	10.51
	(-0.50)	(0.58)	(0.40)	(0.38)		(-0.56)	(0.61)	(0.54)	(0.55)
R^2	0.23	0.59	0.30	0.25	R^2	0.36	0.45	0.37	0.39
HJ-dist	0.34	0.40	0.38	0.35	HJ-dist	0.35	0.39	0.33	0.29
	(0.03)	(0.01)	(0.02)	(0.07)		(0.03)	(0.02)	(0.02)	(0.01)
UK: World Consumption					**UK: Local Consumption**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.019	0.011	0.011	0.013	$\hat{\alpha}$	0.018	0.013	0.015	0.017
	(3.53)	(0.98)	(1.33)	(2.20)		(2.35)	(1.79)	(2.16)	(3.22)
$\hat{\gamma}$	65.68	123.02	41.44	31.45	$\hat{\gamma}$	120.78	53.55	19.70	16.59
	(0.39)	(0.82)	(0.90)	(0.89)		(0.73)	(1.39)	(0.88)	(0.79)
R^2	0.01	0.97	0.88	0.79	R^2	0.89	0.36	0.80	0.72
HJ-dist	0.27	0.20	0.18	0.17	HJ-dist	0.18	0.25	0.21	0.21
	(0.15)	(0.76)	(0.54)	(0.51)		(0.88)	(0.18)	(0.19)	(0.20)
Japan: World Consumption					**Japan: Local Consumption**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.019	0.005	0.016	0.017	$\hat{\alpha}$	0.004	-0.003	0.013	0.014
	(0.72)	(0.41)	(2.07)	(2.17)		(0.30)	(-0.10)	(1.61)	(1.52)
$\hat{\gamma}$	-306.37	34.51	-7.09	-8.14	$\hat{\gamma}$	-71.42	122.31	1.06	1.19
	(-0.53)	(0.45)	(-0.24)	(-0.23)		(-0.46)	(0.44)	(0.08)	(0.09)
R^2	0.53	0.11	0.05	0.05	R^2	0.37	0.52	0.00	0.00
HJ-dist	0.48	0.49	0.50	0.48	HJ-dist	0.48	0.49	0.48	0.47
	(0.00)	(0.00)	(0.00)	(0.00)		(0.00)	(0.01)	(0.02)	(0.02)
Germany: World Consumption					**Germany: Local Consumption**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.007	0.008	0.000	-0.000	$\hat{\alpha}$	0.017	0.018	0.010	0.007
	(0.52)	(0.70)	(0.02)	(-0.01)		(2.40)	(2.12)	(0.56)	(0.31)
$\hat{\gamma}$	516.73	151.77	109.14	96.78	$\hat{\gamma}$	-62.42	-60.28	76.02	67.57
	(1.01)	(0.86)	(0.76)	(0.81)		(-0.88)	(-0.56)	(0.57)	(0.48)
R^2	0.95	0.98	0.89	0.70	R^2	0.70	0.34	0.78	0.44
HJ-dist	0.32	0.23	0.31	0.32	HJ-dist	0.46	0.46	0.48	0.42
	(0.44)	(0.86)	(0.38)	(0.09)		(0.00)	(0.01)	(0.02)	(0.25)
France: World Consumption					**France: Local Consumption**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.009	0.013	0.012	0.011	$\hat{\alpha}$	0.024	0.020	0.017	0.017
	(0.56)	(1.37)	(1.01)	(0.97)		(2.16)	(4.11)	(3.49)	(3.11)
$\hat{\gamma}$	629.80	105.87	58.40	47.28	$\hat{\gamma}$	-233.79	42.51	29.06	26.73
	(0.88)	(0.96)	(1.02)	(1.13)		(-0.71)	(1.41)	(1.38)	(1.74)
R^2	0.63	0.94	0.90	0.83	R^2	0.89	0.90	0.83	0.85
HJ-dist	0.31	0.33	0.37	0.34	HJ-dist	0.35	0.39	0.39	0.37
	(0.10)	(0.06)	(0.02)	(0.01)		(0.56)	(0.03)	(0.01)	(0.01)

Notes: This table reports estimation results for the LR-CCAPM for the within-country cross section of international capital markets (joint pricing of three equity portfolios and five government bond portfolios). Estimation results are presented using discounted long-run world and local consumption growth over horizons $S = 1, 4, 12, 16$. 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 pre-specified weighting matrix. The reported estimation results are for the intercept α, the coefficient of relative risk aversion γ (t-statistic based on the adjustment by Newey and West (1987) with $S+1$ lags in parentheses), the cross-sectional R^2, and the HJ distance (simulation-based p-value in parentheses).

Table 5: Estimation Results LR-CCAPM Across Horizons: Cross-Country Cross Section

Panel A: International Equity Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.026	0.022	0.019	0.020	0.020	0.019	0.016
	(1.35)	(1.93)	(1.55)	(2.03)	(1.93)	(1.68)	(1.50)
$\hat{\gamma}$	-42.85	-15.70	0.41	-4.30	-9.48	-14.04	-18.78
	(-0.32)	(-0.29)	(0.02)	(-0.24)	(-0.57)	(-0.66)	(-0.76)
R^2	0.05	0.06	0.00	0.08	0.30	0.31	0.50
HJ-dist	0.14	0.13	0.13	0.11	0.12	0.13	0.13
	(0.53)	(0.61)	(0.65)	(0.59)	(0.59)	(0.53)	(0.50)

Panel B: International Value/Growth Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.028	0.035	0.025	0.027	0.030	0.032	0.032
	(4.13)	(3.58)	(2.62)	(3.19)	(3.17)	(2.31)	(2.15)
$\hat{\gamma}$	-14.07	-42.42	7.71	1.50	-11.32	-19.01	-25.91
	(-0.17)	(-0.73)	(0.25)	(0.08)	(-0.66)	(-0.82)	(-0.82)
R^2	0.00	0.09	0.00	0.00	0.05	0.06	0.12
HJ-dist	0.50	0.50	0.50	0.50	0.50	0.49	0.50
	(0.00)	(0.00)	(0.01)	(0.02)	(0.06)	(0.05)	(0.02)

Panel C: International Bond Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.013	0.013	0.012	0.011	0.011	0.010	0.010
	(2.39)	(2.80)	(2.64)	(2.81)	(2.32)	(2.38)	(3.01)
$\hat{\gamma}$	-60.36	-30.38	-15.22	3.72	25.69	26.99	18.05
	(-0.40)	(-0.37)	(-0.26)	(0.11)	(0.63)	(0.62)	(0.58)
R^2	0.04	0.06	0.04	0.00	0.17	0.20	0.13
HJ-dist	0.54	0.54	0.54	0.54	0.56	0.58	0.61
	(0.04)	(0.07)	(0.13)	(0.23)	(0.19)	(0.11)	(0.08)

Panel D: All Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.017	0.016	0.014	0.016	0.015	0.014	0.013
	(4.24)	(4.49)	(3.62)	(4.02)	(3.63)	(3.28)	(3.30)
$\hat{\gamma}$	44.60	31.87	39.63	14.58	13.45	15.06	6.93
	(0.53)	(0.48)	(0.80)	(0.52)	(0.61)	(0.76)	(0.23)
R^2	0.02	0.03	0.15	0.05	0.08	0.11	0.02
HJ-dist	0.82	0.81	0.81	0.81	0.83	0.86	0.87
	(0.01)	(0.06)	(0.08)	(0.13)	(0.13)	(0.08)	(0.09)

Notes: This 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 pre-specified 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 and 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 all portfolios jointly (Panel D). All returns are expressed in U.S. dollars.

Table 6: Estimation Results LR-CCAPM Across Horizons: Cross-Country Cross Section (Constant Excluded)

Panel A: International Equity Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	–	–	–	–	–	–	–
	(–)	(–)	(–)	(–)	(–)	(–)	(–)
$\hat{\gamma}$	135.38	97.55	51.29	44.68	37.21	50.75	33.11
	(1.97)	(1.78)	(1.85)	(1.70)	(1.72)	(1.53)	(1.05)
R^2	-0.77	-2.60	-2.55	-7.02	-5.58	-5.44	-2.92
HJ-dist	0.15	0.19	0.18	0.22	0.23	0.22	0.21
	(0.68)	(0.41)	(0.55)	(0.24)	(0.15)	(0.17)	(0.20)

Panel B: International Value/Growth Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	–	–	–	–	–	–	–
	(–)	(–)	(–)	(–)	(–)	(–)	(–)
$\hat{\gamma}$	394.88	139.60	97.88	77.70	64.20	65.18	60.79
	(1.48)	(2.09)	(2.35)	(2.13)	(2.20)	(1.99)	(1.55)
R^2	-2.40	-1.52	-0.74	-1.65	-1.81	-1.07	-1.14
HJ-dist	0.59	0.58	0.59	0.59	0.59	0.58	0.58
	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)

Panel C: International Bond Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	–	–	–	–	–	–	–
	(–)	(–)	(–)	(–)	(–)	(–)	(–)
$\hat{\gamma}$	434.04	179.36	131.96	77.24	110.31	109.90	93.36
	(1.38)	(1.57)	(1.43)	(1.35)	(1.01)	(1.11)	(1.31)
R^2	-2.62	-2.59	-2.91	-2.76	-1.82	-1.85	-2.39
HJ-dist	0.60	0.60	0.59	0.62	0.64	0.68	0.71
	(0.02)	(0.04)	(0.11)	(0.13)	(0.11)	(0.04)	(0.03)

Panel D: All Portfolios							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	–	–	–	–	–	–	–
	(–)	(–)	(–)	(–)	(–)	(–)	(–)
$\hat{\gamma}$	355.83	180.96	114.77	67.93	54.46	53.38	46.54
	(2.36)	(2.34)	(1.94)	(1.67)	(1.79)	(2.08)	(1.59)
R^2	-0.79	-0.64	-0.40	-0.73	-0.79	-0.69	-0.91
HJ-dist	0.84	0.82	0.83	0.84	0.85	0.88	0.89
	(0.01)	(0.05)	(0.08)	(0.13)	(0.12)	(0.07)	(0.07)

Notes: This 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 pre-specified 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 and 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 all portfolios jointly (Panel D). All returns are expressed in U.S. dollars.

Table 7: Decline of the risk aversion Coefficient: Simulation Results

	Case 1			Case 2			Case 3		
	$\Delta = 25$	$\Delta = 50$	$\Delta = 75$	$\Delta = 25$	$\Delta = 50$	$\Delta = 75$	$\Delta = 25$	$\Delta = 50$	$\Delta = 75$
					Int. Equity Portfolios				
A.	0.967	0.918	0.800	0.937	0.841	0.615	0.980	0.920	0.700
B.	0.470	0.446	0.389	0.236	0.212	0.155	0.098	0.092	0.070
					Int. Value/Growth Portfolios				
A.	0.935	0.857	0.650	0.881	0.740	0.361	0.924	0.771	0.333
B.	0.472	0.433	0.328	0.244	0.205	0.100	0.097	0.081	0.035
					Int. Bond Portfolios				
A.	0.969	0.944	0.793	0.944	0.900	0.628	0.969	0.917	0.573
B.	0.469	0.457	0.384	0.254	0.242	0.169	0.093	0.088	0.055
					All Portfolios				
A.	0.956	0.912	0.761	0.912	0.824	0.520	0.987	0.868	0.540
B.	0.480	0.458	0.382	0.228	0.206	0.130	0.075	0.066	0.041

Notes: This table presents simulation results for investigating the decline of the relative risk aversion coefficient in the no constant case. One-step consumption growth rates are bootstrapped (assuming iid consumption growth) to generate 1,000 samples of artificial consumption growth series unrelated to test asset returns by construction. For each of the samples the LR-CCAPM is estimated for horizons $S = 1, ..., 20$ (no constant in the empirical moment function). This table reports the fraction of times a decline of γ greater than 25%, 50% or 75% can be observed in the artificial samples, when, at the same time, $\gamma_{S=1}$ is positive (Case 1), $\gamma_{S=1}$ and $\gamma_{S=20}$ are positive (Case 2) and $\gamma_{S=1,...,20}$ are jointly positive (Case 3). The reported frequencies are computed relative to the samples that fulfill the required property defined by the respective case (A), or relative to the overall number of bootstrap replications (B).

Table 8: Estimation Results LR-CCAPM Across Horizons: Cross-Country Cross Section, U.S. Consumption Growth

I.A. International Stock Markets					**I.B. International Stock Markets**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.015	0.022	0.019	0.017	$\hat{\alpha}$	–	–	–	–
	(0.98)	(1.67)	(1.78)	(1.69)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	37.49	-14.45	-19.06	-22.74	$\hat{\gamma}$	158.89	73.22	31.89	-22.69
	(0.27)	(-0.34)	(-0.72)	(-0.73)		(1.97)	(1.76)	(1.23)	(-0.73)
R^2	0.03	0.13	0.53	0.43	R^2	-0.29	-4.02	-2.88	0.43
HJ-dist	0.14	0.13	0.11	0.12	HJ-dist	0.15	0.19	0.23	0.23
	(0.55)	(0.65)	(0.63)	(0.54)		(0.54)	(0.60)	(0.19)	(0.18)

II.A. International Value/Growth Portfolios					**II.B. International Value/Growth Portfolios**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.029	0.031	0.028	0.025	$\hat{\alpha}$	–	–	–	–
	(3.69)	(2.96)	(2.40)	(2.50)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	-31.23	-17.98	-8.51	2.27	$\hat{\gamma}$	434.62	128.78	90.22	98.36
	(-0.36)	(-0.44)	(-0.30)	(0.08)		(1.76)	(1.91)	(2.48)	(2.24)
R^2	0.01	0.02	0.01	0.00	R^2	-2.08	-1.13	-1.50	-1.81
HJ-dist	0.49	0.48	0.50	0.51	HJ-dist	0.59	0.59	0.59	0.58
	(0.00)	(0.04)	(0.05)	(0.04)		(0.00)	(0.00)	(0.00)	(0.01)

III.A. International Bond Portfolios					**III.B. International Bond Portfolios**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.009	0.011	0.011	0.013	$\hat{\alpha}$	–	–	–	–
	(2.25)	(2.49)	(2.70)	(2.90)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	-113.76	-24.63	45.05	41.20	$\hat{\gamma}$	-389.67	-62.51	191.45	86.23
	(-0.69)	(-0.47)	(1.11)	(1.41)		(-1.54)	(-0.96)	(0.95)	(1.78)
R^2	0.21	0.12	0.25	0.18	R^2	-1.02	-0.14	-2.49	-0.04
HJ-dist	0.55	0.54	0.56	0.58	HJ-dist	0.60	0.60	0.64	0.67
	(0.04)	(0.14)	(0.12)	(0.07)		(0.02)	(0.06)	(0.07)	(0.03)

IV.A. All Portfolios					**IV.B. All Portfolios**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.016	0.015	0.014	0.014	$\hat{\alpha}$	–	–	–	–
	(4.22)	(4.43)	(3.36)	(3.21)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	120.88	38.93	25.95	25.24	$\hat{\gamma}$	244.26	103.64	73.76	71.15
	(1.06)	(0.80)	(0.93)	(1.16)		(1.77)	(1.62)	(2.16)	(3.00)
R^2	0.34	0.20	0.20	0.22	R^2	0.03	-0.37	-0.54	-0.58
HJ-dist	0.82	0.82	0.83	0.86	HJ-dist	0.84	0.83	0.86	0.89
	(0.01)	(0.07)	(0.10)	(0.06)		(0.01)	(0.06)	(0.09)	(0.05)

Notes: This table reports estimation results for the LR-CCAPM for the international asset portfolios. Estimation results are presented using discounted long-run U.S. consumption growth over horizons $S = 1, 4, 12, 16$. 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 pre-specified 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 and 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 all portfolios jointly (Panel D). All returns are expressed in U.S. dollars.

Table 9: Estimation Results LR-CCAPM Across Horizons: Parker and Julliard Specification

I.A. International Stock Markets					**I.B. International Stock Markets**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.028	0.019	0.020	0.020	$\hat{\alpha}$	–	–	–	–
	(1.32)	(1.50)	(2.06)	(1.84)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	-78.43	-0.50	-17.59	-138.27	$\hat{\gamma}$	85.81	27.16	13.87	12.95
	(-0.29)	(-0.02)	(-0.32)	(-0.08)		(3.36)	(3.69)	(4.46)	(4.74)
R^2	0.08	0.00	0.23	0.26	R^2	-0.81	-2.47	-5.48	-5.15
HJ-dist	0.14	0.13	0.12	0.13	HJ-dist	0.16	0.18	0.22	0.21
	(0.80)	(0.85)	(0.86)	(0.79)		(0.82)	(0.75)	(0.33)	(0.35)

II.A. International Value/Growth Portfolios					**II.B. International Value/Growth Portfolios**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.029	0.024	0.031	0.033	$\hat{\alpha}$	–	–	–	–
	(4.00)	(2.34)	(3.21)	(2.59)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	-18.09	9.35	-32.22	115.79	$\hat{\gamma}$	146.28	35.60	15.90	13.29
	(-0.18)	(0.41)	(-0.32)	(0.19)		(3.79)	(6.86)	(6.28)	(6.16)
R^2	0.00	0.01	0.05	0.06	R^2	-2.46	-0.56	-1.56	-0.88
HJ-dist	0.52	0.52	0.51	0.51	HJ-dist	0.60	0.60	0.61	0.59
	(0.00)	(0.02)	(0.09)	(0.08)		(0.00)	(0.00)	(0.01)	(0.01)

III.A. International Bond Portfolios					**III.B. International Bond Portfolios**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.013	0.012	0.011	0.010	$\hat{\alpha}$	–	–	–	–
	(2.44)	(2.62)	(2.30)	(2.29)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	-72.39	-16.96	11.96	11.03	$\hat{\gamma}$	157.40	42.54	19.90	16.73
	(-0.28)	(-0.19)	(1.26)	(1.41)		(3.46)	(3.78)	(4.04)	(5.54)
R^2	0.03	0.03	0.18	0.21	R^2	-2.46	-2.86	-1.83	-1.86
HJ-dist	0.55	0.55	0.56	0.59	HJ-dist	0.60	0.60	0.64	0.68
	(0.23)	(0.27)	(0.25)	(0.15)		(0.12)	(0.21)	(0.14)	(0.06)

IV.A. All Portfolios					**IV.B. All Portfolios**				
Horizon	1	4	12	16	Horizon	1	4	12	16
$\hat{\alpha}$	0.016	0.013	0.015	0.015	$\hat{\alpha}$	–	–	–	–
	(4.02)	(3.40)	(3.63)	(3.37)		(–)	(–)	(–)	(–)
$\hat{\gamma}$	48.05	24.60	8.06	7.65	$\hat{\gamma}$	150.02	40.42	16.42	14.06
	(0.82)	(1.44)	(0.90)	(1.15)		(5.49)	(5.72)	(5.04)	(6.75)
R^2	0.03	0.18	0.09	0.10	R^2	-0.80	-0.34	-0.76	-0.72
HJ-dist	0.83	0.82	0.84	0.87	HJ-dist	0.85	0.84	0.86	0.89
	(0.09)	(0.16)	(0.18)	(0.12)		(0.06)	(0.16)	(0.17)	(0.11)

Notes: This 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 pre-specified 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 and 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 all portfolios jointly (Panel D). All returns are expressed in U.S. dollars.

Table 10: Estimation Results LR-CCAPM Across Horizons: Tests Excluding Japan

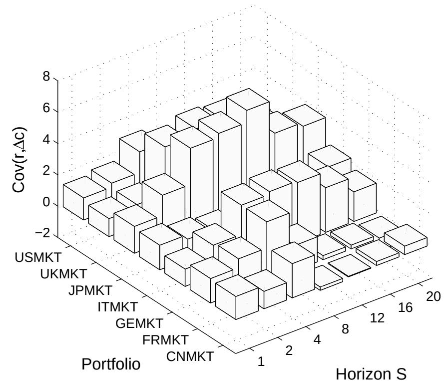
Panel A: International Stock Markets (ex Jap)							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.015	0.014	0.016	0.019	0.020	0.019	0.016
	(0.86)	(0.91)	(1.21)	(2.00)	(1.84)	(1.64)	(1.38)
$\hat{\gamma}$	39.72	40.98	14.96	-0.18	-9.25	-12.83	-23.53
	(0.30)	(0.46)	(0.47)	(-0.01)	(-0.53)	(-0.55)	(-0.88)
R^2	0.04	0.15	0.15	0.00	0.14	0.17	0.43
HJ-dist	0.11	0.12	0.10	0.11	0.11	0.12	0.13
	(0.50)	(0.50)	(0.61)	(0.54)	(0.51)	(0.25)	(0.22)

Panel B: International Value/Growth Portfolios (ex Jap)							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.028	0.031	0.021	0.022	0.023	0.022	0.025
	(3.87)	(3.30)	(1.89)	(2.20)	(2.36)	(2.01)	(2.31)
$\hat{\gamma}$	23.57	-13.30	31.66	25.36	13.98	13.94	-1.02
	(0.27)	(-0.22)	(0.76)	(0.81)	(0.60)	(0.65)	(-0.05)
R^2	0.01	0.01	0.11	0.16	0.04	0.03	0.00
HJ-dist	0.43	0.44	0.45	0.42	0.45	0.45	0.45
	(0.01)	(0.01)	(0.01)	(0.07)	(0.03)	(0.04)	(0.07)

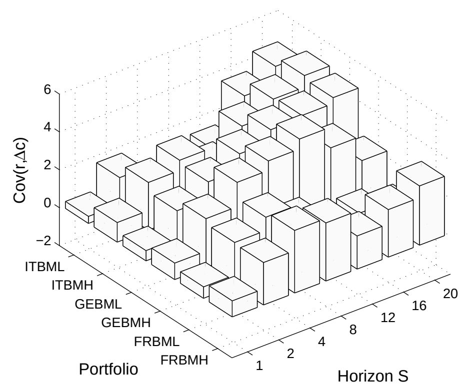
Panel C: International Bond Portfolios (ex Jap)							
Horizon	1	2	4	8	12	16	20
$\hat{\alpha}$	0.012	0.009	0.010	0.009	0.009	0.009	0.009
	(2.35)	(1.35)	(1.70)	(3.37)	(3.00)	(3.13)	(3.50)
$\hat{\gamma}$	17.87	95.11	72.92	51.70	65.39	53.86	52.56
	(0.11)	(0.83)	(0.67)	(0.97)	(0.95)	(0.84)	(0.93)
R^2	0.00	0.14	0.16	0.44	0.59	0.48	0.45
HJ-dist	0.50	0.49	0.50	0.50	0.52	0.53	0.55
	(0.03)	(0.08)	(0.12)	(0.16)	(0.12)	(0.09)	(0.06)

Notes: This 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 pre-specified 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 and West (1987) with $S + 1$ lags in parentheses), the cross-sectional R^2, and the HJ distance (simulation-based p-value in parentheses).

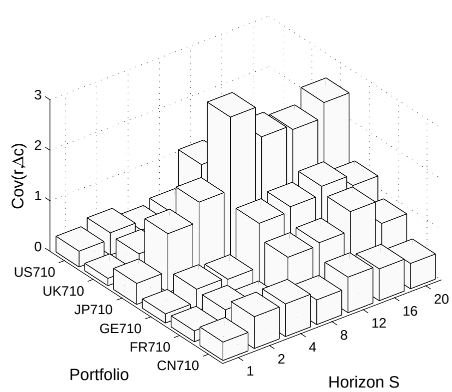
Figure 1: Covariances of Portfolios' Returns with Long-run Consumption Growth



(a) Int. Equity Portfolios



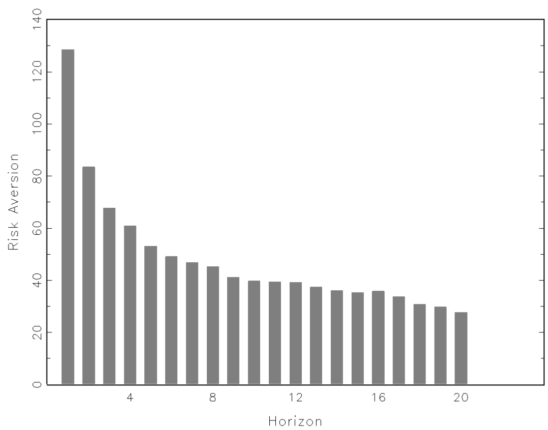
(b) Int. BM Portfolios



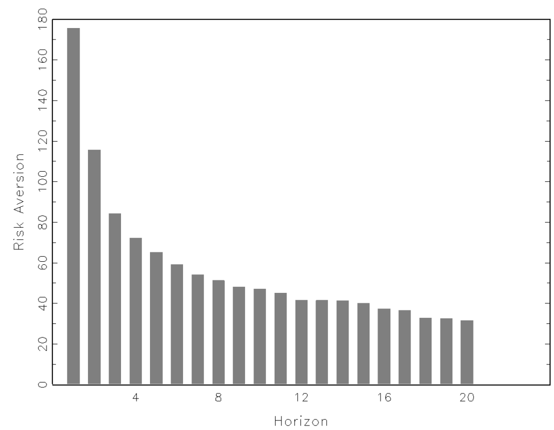
(c) Int. Bond Portfolios, (7-10 years)

Notes: This 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 while subfigure (c) presents covariance plots for the entire set of international bond portfolios (with 7-10 years of maturity).

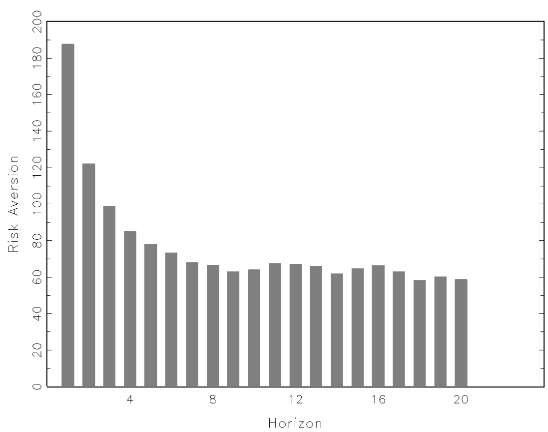
Figure 2: Simulations: Declining Risk Aversion Coefficient



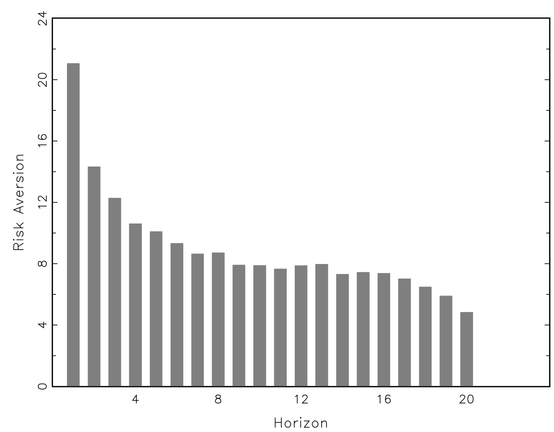
(a) Int. Equity Markets



(b) Int. Value/Growth Portfolios



(c) Int. Bond Portfolios



(d) All Portfolios

Notes: This table of figures illustrates simulation results for investigating the decline of the RRA coefficient in the no constant case. One-step consumption growth rates are bootstrapped (assuming iid consumption growth) to generate 1,000 samples of artificial consumption growth series unrelated to test asset returns by construction. For each of the samples the LR-CCAPM is estimated for horizons $S = 1, ..., 20$ (no constant in the empirical moment function). The mean RRA coefficient (over the 1,000 samples) for which $\gamma_{S=1,...,20}$ are jointly positive (Case 3) is plotted for different horizons.

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