Consumption Heterogeneity and Long Run Risk

Victoria Galsband[1]
University of Zurich

Mathias Hoffmann
University of Zurich & CESifo

December 2007
(Preliminary Version)

[1]University of Zurich, Institute for Empirical Research in Economics, Chair of International Trade and Finance, Zuerichbergstrasse 14, CH-8032 Zurich. E-Mail: victoria.galsband@iew.uzh.ch and mathias.hoffmann@iew.uzh.ch

Abstract

This paper establishes a surprising and robust empirical similarity between short-run heterogeneous consumption and long-term consumption growth risk models. The models not only deliver a similar fit on a given set of portfolios, their actual pricing errors are also highly correlated. In addition, we find that consumption dispersion is a robust predictor of the transitory component in aggregate consumption growth. To interpret these findings, we propose a model in which aggregate uncertainty is a function of idiosyncratic uncertainty and only long-term consumption growth risk is priced. An implication of this being that consumption dispersion is priced *empirically* not because markets are necessarily incomplete but because investors disagree in the short-run about their common long-term consumption prospects.

JEL classification: E21, F30, G15

Keywords: consumption CAPM, idiosyncratic risk, consumption dispersion

1 Introduction

This paper examines the link between two classes of seemingly unrelated empirical asset pricing models. The first class comprises models in which heterogeneity in consumption (measured by the cross-sectional dispersion of consumption growth rates) figures as a pricing factor next to aggregate consumption growth. The second class are models that focus on the role of long-term movements in consumption as a key factor of variation in asset returns.

The theoretical underpinning for the first class of models typically rests on some kind of incompleteness in financial markets that prevents consumers from equating their marginal utilities in equilibrium. Once we aggregate over consumers' first order conditions, this gives rise to consumption dispersion as an additional factor. The second class of models relies on the presence of a small, highly persistent component in aggregate consumption growth. Small changes in the level of this component or in its volatility can have a large impact on the exposure to consumption growth risk in the long-run. If this long-term consumption growth risk is the ultimate risk consumers care about (e.g. because consumption in the short-run is mismeasured or misperceived as in Parker and Julliard (2005)) or if they have a preference for the early resolution of uncertainty (as in Bansal and Yaron (2004)), long-term consumption growth risk should be a much more potent pricing factor than short-term consumption growth – an implication that is generally confirmed in the data.

Our contribution here is to document a surprising and robust empirical similarity between versions of the C-CAPM with heterogeneous consumption (to which we refer as HC-CAPM) and those with long-term consumption growth risk (that we call LRC-CAPM). First, consumption dispersion is highly correlated with and a robust predictor of the transitory component in aggregate consumption growth. Secondly, the HC-CAPM and the LRC-CAPM not only perform similarly well on a wide range of test assets at different time horizons, they actually generate almost the same expected returns and highly correlated pricing errors. This suggests that the two models are just two different and almost equivalent empirical incarnations of the same underlying theoretical mechanism.

We capture this idea in a simple theoretical model in which agents care only about permanent shocks to consumption so that long-term consumption growth risk is the relevant pricing factor. In the short-run, however, agents have heterogeneous perceptions about long-term consumption growth. The uncertainty about ultimate consumption growth goes up in downturns and is reflected in increased heterogeneity of observed individual consumption decisions. Even though this is a model of long-term consumption growth risk, it implies that consumption dispersion and current average consumption growth should jointly capture the information embodied in long-term growth prospects, making them valid stand-ins for long-run consumption growth as empirical pricing factors.

An important feature of our setup is that it gives rise to consumption dispersion as an empirical pricing factor even if markets are complete. The reason why heterogeneity is – empirically – priced here is just that informational asymmetries are high when aggregate consumption growth is low relative to its long-run path. Hence consumption dispersion conditioned on current consumption growth becomes a sufficient statistics for the transitory component in aggregate consumption growth.

Our model has a couple of ancillary implications that we explore further. Specifically, it allows us to write the LRC-CAPM as an economically motivated two-beta model. In this setup, we show that the documented similarity between HC- and LRC-CAPMs might be due to a strong negative relation between long-run consumption growth and dispersion betas. We find that assets with high exposure to ultimate consumption risk do indeed deliver a low exposure to idiosyncratic consumption risk.

The plan for the remainder of the paper is as follows. We present our empirical HC- and LRC-CAPMs in the next section, providing evidence of the surprising similarity in their empirical performance. Section three presents the theoretical model that we put forward to explain this similarity. We explore the implications of our model further in section four and conclude in section five.

2 The two empirical models

The two pricing models we confront build upon the key insight of standard canonical C-CAPM that differences in expected returns can be explained via asset's exposure to consumption risk.

Instead of studying contemporaneous comovements of returns with consumption, long-term risk models examine whether assets can be priced by their exposure to long-run consumption risk. This approach maintains the assumption that it is the low-frequency component of consumption path that is actually priced. One reason for that being mismeasurement of high frequency data. Another intuitive explanation is that consumption is slow to adjust. In particular, the LRC-CAPM considers the intertemporal marginal rate of substitution (IMRS) in the permanent component of consumption, c_t^P, as a valid stochastic discount factor (SDF). Specifically, under time-separable utility with constant relative risk aversion (TS-CRRA) the log version of the long-run consumption based pricing kernel is given by

$$m_{t+1}^{LRC} = 1 - b_{\Delta c} \triangle c_{t+1}^P, \tag{1}$$

where the constant is normalized to 1, factor loading $b_{\Delta c} \geq 0$ governs the risk aversion coefficient and lowercase letters refer to logarithms of corresponding capital letters. From equation (1) it follows that the only relevant fundamental pricing factor is the true permanent world consumption growth[1], denoted by $\triangle c_{t+1}^P$.

Short-run heterogeneous consumption models amplify the real SDF by letting it account for uninsured idiosyncratic consumption risk. With TS-CRRA specification, the HC-CAPM is derived by taking a second order Taylor expan-

[1] For estimation, we will use $(c_{t+s} - c_t)$ as a proxy for $\triangle c_{t+1}^P(s)$, in the spirit of Julliard and Parker (2005), where s denotes the horizon in quarters over which the consumption response is studied.

sion of IMRS between t and $t+1$:

$$M_{t+1}^{HC} = \delta \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma} \exp \left[\frac{\gamma(\gamma+1)}{2} Var \left(\frac{C_{k,t+1}/C_{t+1}}{C_{k,t}/C_t} \right) \right], \tag{2}$$

In equation (2) above, C_t and $C_{k,t}$ are world average and country k consumption expenditure per capita at time t, respectively, γ denotes the coefficient of RRA, and the world consumption dispersion is measured as the variance of cross-country distribution of relative consumption growth rates. To simplify the notation, we henceforth abbreviate the world consumption dispersion across K countries as $\sigma_{K,t+1}^2 = Var_K \left(\frac{C_{k,t+1}/C_{t+1}}{C_{k,t}/C_t} \right)$. Models capturing the cross-sectional skewness term provide a natural framework within which the effects of heterogeneity can be studied. Interestingly, there are several ways of going about the sources of heterogeneity. While Constantinides and Duffie (1996) initially generate the $\sigma_{K,t+1}^2$-term as cross-sectional variance of the $N(0,1)$ highly persistent idiosyncratic shocks arising due to incomplete markets structure, Bansal and Yaron (2004) exploit the variance term as the relevant state variable proxying the time-varying economic uncertainty. Some other authors, i.e. De Santis (2005), construct models with conditional variance varying slowly over time and reminding essentially of habit persistence setups.

In our study, we interpret the $\sigma_{K,t+1}^2$-component as a dispersion factor that comes to life as a result of informational heterogeneity reflected in individual misperceptions about possible realizations of long-term consumption growth. Furthermore, we model cross-sectional heterogeneity in our setting, as a direct determinant of time varying aggregate uncertainty. Analogously to (1), the log-linearized SDF of the HC-CAPM can be compactly summarized as:

$$m_{t+1}^{HC} = 1 - b_{\triangle c} \triangle c_{t+1} + b_{\sigma} \sigma_{K,t+1}^2, \tag{3}$$

where the vector of factors now consists of short-run growth prospects and heterogeneous perceptions about future economy. According to (3), the innovation in m_{t+1}^{HC} is driven by the innovations in short-run $\triangle c_{t+1}$ and $\sigma_{K,t+1}^2$, and the risk premium for any traded asset is determined by covariation of returns with the innovations in m_{t+1}^{HC}.

Using the standard asset pricing restriction for gross real rate of return[2] to market portfolio of country k, R_{t+1}^{k}, it follows, that

$$E_t \left[\exp \left(1 - b_{\Delta c} \triangle c_{t+1}^{P} \left(s \right) + r_{t+1}^{k} \right) \right] = 1 \tag{4}$$

and

$$E_t \left[\exp \left(1 - b_{\Delta c} \triangle c_{t+1} + b_{\sigma} \sigma_{K,t+1}^{2} + r_{t+1}^{k} \right) \right] = 1, \tag{5}$$

where $E_t \left[\cdot \right]$ denotes the expectation operator based on information available at time t and $\log \left(R_{t+1}^{k} \right) \equiv r_{t+1}^{k}$.

In the following subsection, we give a brief description of the data set we use to compare the empirical performance of the two models presented.

2.1 Data

The data set is quarterly, the sampling interval covers a period from 1973Q1 to 1996Q1. The global international equity markets considered in this article are comprised of 8 industrialized countries (G-7 plus Switzerland) and the world index. The data set on international equity indices includes quarterly *Morgan Stanley Capital International (MSCI)* stock market data covering the post-Bretton-Woods period. The data is freely available at www.mscibarra.com. We calculate quarterly returns from end-of-quarter index level denominated in U.S. dollars. Excess returns are constructed by subtracting the return on a three-month US Treasury bill from these returns. To study the effect of consumption dispersion on asset prices we deliberately switch off the effect of the foreign exchange markets by choosing unitary (dollar-denominated) as opposed to national currency denominated returns.

The macroeconomic data on seasonally adjusted aggregate consumption, to-

[2] In the empirical work, we abstract from exchange rate movements. The focus of our research lies on international unitary (dollar-denominated) stock returns. A precise specification of equity markets data is provided in Section 2.1.

tal population, and the gross domestic product come from the *National Accounts*. Our sample contains data on real (chain-weighted) personal consumption expenditures on nondurable goods of Canada, France, Germany, Italy, Japan, Switzerland, United Kingdom, and the Unites States, with the latter being the domestic country. To obtain national per capita consumption in local currency[3], the aggregate real consumption for each country is deflated by the estimates of quarterly population in the economy. These estimates are obtained by linear interpolation of annual population data. Logarithmic growth rates of real per capita variables are calculated as first differences of natural log of per capita deflated level values. The world consumption growth rate measure is constructed as the GDP-weighted average of single countries' consumption growth rates. Apparently, the GDP weights attach a higher value to consumption growth of countries with higher GDP. Since consumption growth, unlike consumption levels, is unitless, average world consumption growth is a GDP-weighted average of real per capita consumption growth rates in local currency of the 8 countries under consideration.

To arrive at cross-sectional variance measure of world consumption growth we calculate the variance of the log changes in national real per capita consumption growth rates in local currency units of individual countries at each point in time.

In line with existing literature, we make the standard "end-of-period" timing assumption that consumption during quarter t occurs at the end of the quarter, so that $\triangle c_{t+1}$ is calculated using consumption in period $t+1$ relative to t. Furthermore, we use a two-dimensional timing convention in the article: By s we denote the horizon in quarters over which the consumption response is studied. We employ index h to denote the horizon in quarterly frequency over which the returns are cumulated.

[3]Sarkissian (2003) gives the rationale for using consumption data in local currency units as opposed to consumption data expressed in U.S. dollars.

2.2 Estimation methodology

To evaluate the economical importance of aggregate consumption and heterogeneity in individual consumption decisions, we use Hansen's (1982) generalized method of moments[4] (GMM). In particular, we test the conditional orthogonality equations by using the SDF representations[5] in (1) and (3). The vector of constant parameters is chosen in such a way that the pattern of returns fulfills best the asset pricing conditions in (4) and (5).

Following Hansen and Jagannathan (1997), the models are estimated by two-stage GMM. To obtain the estimates of $b_{\Delta c}$ and b_{σ}, we employ the identity matrix as the weighting matrix in the first stage. As suggested by Hansen and Singleton (1982), we use the inverse of the spectral density as the optimal weighting matrix in the second stage. The instrument vector is composed of a constant and lagged values of world returns. Given 1(2) parameters of interest and 18 orthogonality conditions, this implies 17(16) degrees of freedom. We begin the estimation by using the prespecified weighting matrix. Another part of our testing strategy relies on efficient GMM, where the weighting matrix is given by the sample covariance matrix of the orthogonality conditions[6]. Our first set of results is presented in Table 1. Each row of Table 1 represents a set of estimation results for the standard C-, LRC-, and HC-CAPM specifications for different return horizons h. Coefficient estimates are reported in the first row. The second row gives the Newey-West corrected t-statistics. Hansen's J-test statistic of overidentifying restrictions and corresponding p-values are listed in the last column. Panel A of the table reports the results for the standard C-CAPM. Contemporaneous consumption risk implies implausibly large levels of γ. The point estimate of risk aversion required to rationalize the cross section of

[4]This tesing procedure has been implemented in a broad variety of recent empirical studies in asset pricing (Cochrane (2001), Heaton and Lucas (1996), Jacobs and Wang (2004), etc.).

[5]The cross section of asset returns can be analyzed directly, i.e. using the SDF from (2). However, the existence of local optima usually complicates the optimization process. To circumvent estimation hurdles in highly nonlinear relations, we test (1) and (3), assuming that the pricing kernels depend in a linear way on the factors.

[6]Because the two-stage GMM is asymptotically efficient, it forms a natural starting point to explore the statistical significance of the factor loadings.

international returns at quarterly frequency is roaming about 190. Despite the low fit of the model, contemporaneous consumption risk is typically statistically significant using both estimation methodologies. Panel B indicates in a forceful way that low frequency consumption data provides a better statistic for pricing assets vis-à-vis high frequency consumption data. In line with the existing literature, the point estimates of $b_{\Delta c}$ are always lower for the ultimate risk model compared to standard C-CAPM. However, despite the economic significance of long term consumption risk and far lower CRRA estimates, the data reject that the single long-run consumption factor is the only determinant of expected returns. Using efficiently reweighted moments lowers CRRA by about 10% at all horizons. Otherwise, our conclusions remain qualitatively similar to those using prespecified first stage GMM. The last set of results is indicative of economic and statistical importance of the dispersion factor. Taking account of heterogeneity in consumption decisions in Panel C lowers the γ estimates from Panel A by close to 50%. Interestingly, the LRC-CAPM appears to deliver $b_{\Delta c}$ estimates of even more plausible order of magnitude. Despite strongly significant estimates, the overall fit of the models is poor. All overidentification test statistics indicate rejection of the null hypothesis at conventional levels of significance. Yet, this result may not necessarily be surprising. Julliard and Parker (2005) and Jacobs and Wang (2004), among others obtain comparable outcomes.

To conclude, Table 1 clearly demonstrates a high significance of the main risk factors under consideration. In all specifications that we have analyzed, ultimate consumption risk and short-run dispersion factor are estimated with a right sign, and both are statistically significant. Our results suggest that the performance of model specifications in Panels A-C is not very different in terms of significance and the J-statistics. In contrast to comparisons based on chi-square statistics, such as Hansen statistic, the Hansen-Jagannathan distance (Table 2) has several desirable properties. First of all, it does not reward variability of SDF. Secondly, the weighting matrix remains the same across various pricing models, which makes it possible to compare the performances among competitive SDFs by the relative values of the distances. Estimating the HJ-distance reveals the first stylized fact, namely very close estimates for the models with heterogeneity and long run consumption risk.

2.3 The puzzle

Using standard linear regression methods, we now examine the empirical support of the dichotomy between cross-sectional time-varying heterogeneity and long-run changes in international consumption growth rates. In what follows, we compare the performance of the LRC-CAPM and two-factor HC-CAPM with dispersion as a second pricing factor.

2.4 Cross-sectional perspective

To examine the extent to which the cross-section of returns predicted by both models can explain each other, we regress fitted returns in HC-CAPM on fitted returns in LRC-CAPM. We include an intercept that allows all average returns predicted by the two models to differ by a common amount. At horizons of one to twelve quarters, the point estimates of the intercept term remain roaming around zero. Note that the slope estimates are very close to unity, reaching the value of 1.07 for biennial returns. The results suggest that to a large extent both models may be capturing the same consumption related world-wide risks. Two main points of Table 3 are (1) that the performance of short-run model that accounts for dispersion in consumer expectations rivals the single-factor long-term consumption risk model; and (2) that there is a tendency for the models to fit each other better as the holding period return increases. As we increase the horizon, there is an improvement in the extent to which the models explain each other, fitting 80-92 percent at different horizons.

Figure 3 gives the plot of average returns predicted by the two models. Note that the points are roughly evenly distributed around the 45-degree line[7]. This

[7] At short time horizons, the ultimate risk tends to overestimate returns relative to the two-factor HC-CAPM with six points lying above the 45-degree line.

finding is robust for different time horizons. The degree of correlation between the predicted returns jumps from 72 percent for quarterly returns up to slightly more than 90 percent for biennial returns.

2.5 Individual asset level

Our finding seems to be qualitatively unaffected once we go down to individual asset level. In a simple exercise, we run time-series regressions of the following form:

$$r^k_{t+h} = \alpha^k_h + \beta^k_h \Delta c^P_{t+h} + \varepsilon^k_{t+h},$$

in the case of the LRC-CAPM and

$$r^k_{t+h} = \alpha^k_h + \beta^k_{1,h} \Delta c_{t+h} + \beta^k_{2,h} \sigma^2_{K,t+h} + \varepsilon^k_{t+h}$$

for the HC-CAPM. In both regressions above, r^k_{t+h} is return of country k at time $t + h$. As has been shown in the literature of the last ten years, one should treat long-horizon regressions with a lot of caution. Using asymptotic methods, Valkanov (2003) shows that the t-statistics of the OLS-estimators in long-horizon regressions do not converge to well-defined distributions. Moreover, in certain cases the estimators are biased and the R^2 statistic is no longer an adequate measure of the goodness of fit. To conduct inference using long-horizon regressions we rely on solution provided by Valkanov (2003). Pooled regressions in Tables 4 and 5 yield highly significant parameter estimates at virtually all differencing horizons. The time series variation in international risk premia is explained best at horizons of 12 quarters, yielding adjusted R^2 close to 20%. Interestingly, estimates on short-term consumption risk do not have an expected sign in Panel A of Table 5. In line with Parker (2003), we find evidence that contemporaneous consumption risk is negatively related to time variation in expected returns. The estimates become again positive for single country time series regressions at longer horizons. Table 5 reveals that fluctuating economic uncertainty directly affects the cross-sectional properties of international returns. A rise in economic uncertainty leads to a fall in asset prices. This result is robust for factor cumulation over 1 to 12 quarters. Moreover, it holds true with some minor exceptions once we switch from pooled to single country time series regressions in Panel B. Leaving the risk-free rate out of regressions does not alter the results qualitatively. Both regression types in Tables 4 and 5 deliver a similar fit for different holding period returns. Moreover, both models generate highly correlated errors at different return horizons. To visualize this fact, Figure 4 graphs the term-structure of mean squared errors for each country, implied by two models at horizons between 1 and 15 quarters. The pricing errors are highly correlated for all countries under consideration with correlation coefficients varying from 63% up to 92%. Hence, also at the individual asset level, the two models perform about equally well in explaining the time-series variation in returns.

2.6 Pricing consumption and heterogeneity risks

We present the results from second stage of the Fama-MacBeth (1973) method along with the Shanken corrected t-statistics in Table 6. In the second stage, the full-sample beta estimates from the first stage are used as explanatory variables. At each time period, we run a cross-sectional regression of returns on the betas that were estimated in the first pass. Risk prices are then backed out as the average of cross-sectional regression estimates in the second stage of the procedure, and the standard errors are calculated from the time series standard deviation of the estimates[8].

A correct specification implies an intercept term equal to the risk-free rate. That is, assets bearing no consumption associated risk should earn a risk premium amounting to the prevailing risk-free rate in the economy. Note that the

[8] Cochrane (2001) shows that Fama-MacBeth standard errors do not include corrections for the fact that the betas are also estimated. The resulting t-statistics are in fact corrected for cross-sectional correlation but not for time-series correlation in the residuals. Shanken (1992) provides a correction for t-statistics comprising of a multiplicative and an additive term.

intercept estimated in Table 6 is about 0.10-0.13% per year. This pattern of numbers has also been observed for instance by Jagannathan and Wang (2007).

Panel A of Table 6 shows a parameterization implied by standard C-CAPM. Despite its significance, the point estimate of structural parameter of interest is economically very low. Panel C of Table 6 demonstrates some improvement in explanatory power gained by considering the ultimate risk to consumption within a LRC-CAPM. Even though the increase in R^2 is small, there is a dramatic improvement in the magnitude of coefficient estimate. After taking sampling errors into account, the slope remains significantly positive, consistent with the view that consumption risk carries a positive risk premium. The estimated return per unit ultimate risk is about 15 basis points per annum.

Panel B documents a better performance of HC-CAPM vis-à-vis standard C-CAPM due to cross-sectional dispersion factor additionally entering the regression. Inclusion of heterogeneity into the short-term pricing kernel provides a huge extra benefit. More than 50 percent of the variation in expected returns is now explained by the short-run model. Furthermore, in standard Fama-MacBeth cross-sectional regressions, the two-factor model with economic uncertainty factor provides more explanatory power than the model of long-term consumption risk. The estimated value of the consumption risk premium of 0.02% p.a. is as anticipated positive and does not differ much from the estimate of 0.013% p.a. using the standard C-CAPM. However, the estimated risk premium for the dispersion factor is found to be negative, lying by -1.92% per year. The negative estimates are supported by our beta decomposition framework below (compare also Figure 6). There we show that α_σ is indeed a negative function of cross-sectional variance of consumption dispersion.

In Panel D we quantify the impact of heterogeneity, given long-run growth prospects of the economy. The t-statistic of α_σ shows that dispersion beta is no longer a significant determinant of the cross section of average returns, once long-run consumption growth has been explicitly accounted for. Dispersion beta neither seems to have an important effect on the marginal predictive power of the long-term consumption risk nor does it substantially improve the overall fit. A possible reason is that short run heterogeneity factor bears a similar information set with respect to expected returns as ultimate consumption risk does. Therefore, the idiosyncratic economic uncertainty seems to add noise rather than signal, once the real ultimate consumption growth is already included as a regressor. This finding is largely consistent with our maintained hypothesis of predicting potential of short-run dispersion for the transitory component in aggregate consumption growth. We conjecture that short-run dispersion seems to contain information about long-run ultimate growth prospects of the economy on average. Given that a consumer has already included the long-run mean of consumption growth into his information set, the dispersion factor brings essentially no additional information and looses its significance as a risk factor. Hence, this finding appears to reinforce the insight of international consumption dispersion being informative of world average consumption growth in the long run.

Our overall results suggest that whereas the dispersion factor is priced at shorter horizons, markets price consumption growth not only for the short but also the very long run. However, aggregate consumption risk alone prices the expected returns rather poorly suggesting that issues of heterogeneity are important elements of the story. Fama-MacBeth method supports our finding that a two-factor model with risks related to varying short-run growth prospects and fluctuating economic uncertainty and a single-factor ultimate consumption growth model basically seem to do the same job (see Figure 2).

Further we elaborate the major claim of our paper that the main force driving the similarity of both models is precisely the deep link between the dispersion series and the transitory component in long-term consumption growth. Short run heterogeneity in consumption decisions coupled with current consumption data seems to bear a similar information set with respect to expected returns as ultimate consumption risk does. Following this line of argumentation, time-varying uncertainty appears to play a two-fold role in our analysis: First, it is informative of the near-term properties of cross-sectional consumption growth volatility. Secondly, it seems to contain an information set about what long-run consumption growth might be. We explore this feature further in ancillary implications section.

3 A simple theoretical explanation

We now present a simple theoretical framework that rationalizes the stylized fact presented in the previous section. We conjecture that for each country k, consumption growth is the sum of a common priced factor which constitutes ultimate consumption growth risk, and a temporary idiosyncratic component, so that

$$\Delta c_t^k = \Delta c_t^P + \Delta c_t^{Tk}, \tag{6}$$

where Δc_t^P is the low-frequency component of consumption growth. It is this component which is priced in our model, such that the Euler equation is the same for all countries. Using CRRA-setup, gross returns on market portfolio in country k, denoted by R_{t+1}^k, should therefore obey:

$$\boldsymbol{E}_t((1 + \Delta c_{t+1}^P)^{-\gamma} R_{t+1}^k) = 1. \tag{7}$$

There could be a number of reasons why only a subcomponent of actual consumption growth enters the pricing relation. The simplest explanation would be that consumption is mismeasured or misperceived by agents in the short run. Also, consumption data that would allow cross-regional or international comparisons, would generally include expenditure on durable goods. However, the economically relevant concept of consumption should be the consumption of non-durables plus the unobserved service flow from the stock of durables (Campbell and Mankiw (1989)). Since durables consumption tends to be rather volatile, focusing on low-frequency movements in consumption growth may effectively allow to purge consumption from these effects. A variant of this argument is that consumers have more discretion over expenditure on luxury goods than on essential goods.

Another possibility is that consumption is influenced by high-frequency preference shocks that have, however, no influence on household portfolio decisions which are taken only in larger intervals. In this mould, Jagannathan and Wang (2007) argue that annual growth rates based on fourth quarter consumption decisions are much more correlated with stock returns than are annual growth rates from other quarters, because investors are 'lazy' and take most of their major portfolio decisions at the end of the year.

In our model the transitory component of consumption varies across countries. Specifically, we assume that Δc_t^{Tk} can be written as

$$\Delta c_t^{Tk} = \sigma_{K,t-1}(\mu_t + \eta_t^k), \tag{8}$$

where $\sigma^2_{K,t-1}$ denotes the cross-sectional variance of growth rates across K countries as in (3). Moreover μ_t is the transitory component of consumption growth that is common to all countries, whereas η_t^k is a country-specific component with cross-sectional mean zero and unit variance:

$$\begin{aligned} \boldsymbol{E}_K\left(\eta_t^k\right) &= 0 \\ var_K\left(\eta_t^k\right) &= 1. \end{aligned}$$

In this setup, it is easily verified that aggregate consumption growth is given by the cross-sectional average

$$\overline{\Delta c}_t = \Delta c_t^P + \sigma_{K,t-1}\mu_t \tag{9}$$

and $var_{K,t-1}(\Delta c_t^k) = \sigma^2_{K,t-1}$. Now let

$$\varepsilon_t = \overline{\Delta c}_t - \boldsymbol{E}_{t-1}(\overline{\Delta c}_t) \tag{10}$$

be the shock to aggregate consumption growth. Then

$$\varepsilon_t = \Delta c_t^P - \mathbf{E}_{t-1}(\Delta c_t^P) + \sigma_{K,t-1}(\mu_t - \mathbf{E}_{t-1}(\mu_t)) \tag{11}$$

and the conditional variance of aggregate consumption becomes

$$\omega_{t-1}^2 = var_{t-1}(\Delta c_t^P) + 2\sigma_{K,t-1}cov_{t-1}(\Delta c_t^P, \mu_t) + \sigma_{K,t-1}^2 var_{t-1}(\mu_t). \tag{12}$$

Following Bansal and Yaron (2004), we abstract from time-variation in the conditional variance of permanent consumption growth. Also we assume that the correlation ρ between Δc_t^P and μ_t is zero so that (12) simplifies to

$$\omega_{t-1}^2 = \sigma_{K,t-1}^2 var_{t-1}(\mu_t) + const. \tag{13}$$

Equation (13) captures a key feature of our model: It decomposes aggregate uncertainty (as measured by the volatility in aggregate consumption) into a common component ($var_{t-1}(\mu_t)$) and an idiosyncratic component ($\sigma_{K,t-1}^2$). Given the common component, aggregate uncertainty is a function of heterogeneity in individual consumption decisions. The economic interpretation is straightforward. If the path of the aggregate economy is highly uncertain, this can have two causes. The first is that there is just a lot of uncertainty about the variability in common shocks. The second, on which we focus here, is that people have very different perceptions of the economy, which leads to a lot of heterogeneity in observed decisions.

An important feature of the data is that aggregate consumption growth is negatively related to the cross-sectional dispersion of consumption growth. In particular, we will document below that this stylized fact is largely due to a negative correlation between the transitory part of consumption growth and dispersion. We therefore capture the dynamics of the common transitory component by postulating that shocks to dispersion

$$\xi_{t+1} = \sigma^2_{K,t+1} - E_t(\sigma^2_{K.t+1}) \tag{14}$$

are perfectly negatively correlated with shocks to the common transitory component:

$$\mu_{t+1} - \boldsymbol{E}_t(\mu_{t+1}) = -\delta\xi_{t+1} \text{ where } \delta > 0. \tag{15}$$

We now work out the pricing implications from this model. Using the decomposition from (9) above, we write

$$\Delta c^P_{t+1} = \overline{\Delta c}_{t+1} - \sigma_{K,t}\mu_{t+1} \tag{16}$$

so that the Euler equation (7) becomes

$$\boldsymbol{E}_t((1 + \overline{\Delta c}_{t+1} - \sigma_{K,t}\mu_{t+1})^{-\gamma} R^k_{t+1}) = 1. \tag{17}$$

Then it should approximately hold true that

$$\boldsymbol{E}_t(\exp\left[-\gamma(\overline{\Delta c}_{t+1} - \sigma_{K,t}\mu_{t+1})\right] \exp(r^k_{t+1})) = 1. \tag{18}$$

Assuming that all random variables in this expression are jointly log-normal and using the fact that the risk-free rate in our model is given by

$$r^f_t = \gamma \boldsymbol{E}_t(\overline{\Delta c}_{t+1} - \sigma_{K,t}\mu_{t+1}) - \frac{\gamma^2}{2}\boldsymbol{var}_t(\overline{\Delta c}_{t+1} - \sigma_{K,t}\mu_{t+1}) \tag{19}$$

we can write the excess return on the market portfolio of country k as

$$\boldsymbol{E}_t(r^k_{t+1} - r^f_t) = \gamma \mathbf{cov}_t(\overline{\Delta c}_{t+1}, r^k_{t+1}) - \gamma\sigma_{K,t}\mathbf{cov}_t(-\delta\xi_{t+1}, r^k_{t+1}) + \kappa \tag{20}$$

and $\kappa = -\frac{1}{2}\mathbf{var}_t(r^k_{t+1}) - \gamma\sigma_{K,t}\mathbf{cov}_t(\boldsymbol{E}_t\left(\mu_{t+1}\right), r^k_{t+1})$.

Hence, our model of long-run consumption risk can be written as a two-factor model in aggregate consumption growth and dispersion[9]. In this setup, the dispersion factor is, however, just a correction for the difference between aggregate consumption growth and long-term consumption growth, not an indicator of the fact that idiosyncratic risk is priced.

We now flesh out some implications from this model.

4 Ancillary implications

The theoretical explanation we gave for the empirical similarity between the LRC-CAPM and the HC-CAPM implies a decomposition of aggregate consumption risk into a common permanent and a transitory idiosyncratic components. In this section, we provide empirical support for this decomposition.

4.1 Beta decomposition framework

We turn to the implications of beta representation first. Using (16) we can write an asset's k exposure to long-term consumption growth risk as

$$Cov\left(r_{t+1}^{k}, \triangle c_{t+1}^{P}\right) = Cov\left(r_{t+1}^{k}, \overline{\Delta c}_{t+1} - \overline{\Delta c}_{t+1}^{T}\right). \tag{21}$$

The expected returns on portfolio k are therefore governed by the covariance with world aggregate consumption growth, on the one hand, and idiosyncratic consumption growth, on the other hand. Using the link between the transitory part of aggregate consumption and the cross-sectional dispersion implied by our model, i.e. $\overline{\Delta c}_{t+1}^{T} = \sigma_{K,t}\mu_{t+1}$, it is now easy to demonstrate that the 'true', i.e. long-term beta can be written as:

$$\beta_{k}^{P} = \beta_{k,\overline{\triangle c}} \frac{Var\left(\overline{\Delta c}_{t+1}\right)}{Var\left(\triangle c_{t+1}^{P}\right)} - \beta_{k,\sigma} \frac{Var\left(\sigma_{K,t}\mu_{t+1}\right)}{Var\left(\triangle c_{t+1}^{P}\right)}, \tag{22}$$

[9] Note that the second factor is actually scaled with lagged dispersion. We explore this implication below, but generally implement (20) without conditioning information.

where

$$\beta_{k,\overline{\triangle c}} \equiv \frac{Cov\left(r^k_{t+1}, \overline{\Delta c}_{t+1}\right)}{Var\left(\overline{\Delta c}_{t+1}\right)} \tag{23}$$

is the asset's beta with respect to aggregate consumption growth risk and

$$\beta_{k,\sigma} \equiv \frac{Cov\left(r^k_{t+1},\ \sigma_{K,t}\mu_{t+1}\right)}{Var\left(\sigma_{K,t}\mu_{t+1}\right)} \tag{24}$$

denotes its exposure to consumption dispersion. Given fixed $\beta_{k,\overline{\triangle c}}$, high exposure to ultimate consumption growth implies low exposure to idiosyncratic consumption risk and vice versa. Our data support the view that there is very little variation in portfolio's systematic exposure to aggregate consumption growth risk, i.e. in $\beta_{k,\overline{\triangle c}}$ (compare for instance Figure 1, Panel A). We find that a regression of the form

$$\beta^P_k = \alpha_0 + \alpha_\sigma \beta_{k,\sigma}, \quad \text{with} \qquad \alpha_\sigma < 0$$

provides a quite reasonable fit for the relation between the long-run and dispersion betas and that it gives a significantly negative slope. Figure 8 provides an optical impression of this link.

4.2 Predictive power of dispersion factor

This section asks whether idiosyncratic risk proxied by the short-run cross-sectional dispersion in international consumption growth rates can explain the variation in transitory component of average world consumption growth rate. To get at this issue we examine the long-horizon forecastability of dispersion for the relative consumption growth. We measure the transitory component in consumption by consumption growth spread as follows. First, we calculate the equivalent quarterly average corresponding to the ultimate consumption growth over an s-quarter period. We subtract it then from first quarter aggregate growth rate. Finally, we regress the obtained transitory component on the short-run dispersion factor at the end of first quarter. Thus the OLS regressions we run are of the following type:

$$\triangle c_{t+1}^{T}(s) = \beta_0 + \beta_s \sigma_{K,t+1}^2 + \varepsilon_{t+s}.$$

Long-horizon regressions suggest a strong ability of heterogeneity factor to forecast the transitory variation in aggregate consumption growth rate. The slope coefficient increases in absolute terms from 7.15 to 18.47 when the horizon spread lengthens from 1 to 12 quarters[10]. Table 7 also shows that all t-statistics increase in the time horizon. Several other features bear noting. First, the results are consistent with the basic intuition that if long-run growth rates are predictable by a slow-moving variable, the predictability should build up with the horizon: The obtained R^2's more than triple ranging from 3.8% for one-quarter spread to slightly more than 13% when s =13. However, despite the strong relationship between the transitory part of consumption growth risk and short-run dispersion in growth rates, heterogeneity motive does not provide a complete accounting for relative ultimate consumption risk.

To present another perspective on the results, we look at the evolution of the correlation pattern between the transitory component in growth and dispersion over time. Focusing on partial correlations, Table 8 clearly supports a strong economic and significant statistic correlation between consumption spread and cross-country consumption dispersion: A decline in cross-sectional consumption volatility goes along with a higher wedge between aggregate and permanent consumption growth rates. Notice that the correlation of two series increases over the first 2 years stagnating then at a level of about 37%. A visual perspective on the comovement pattern is provided in Figure 9. Apparently, the figure demonstrates a striking negative comovement (with $\rho = -0.475$) of transitory component in consumption growth and short-run consumption dispersion factor.

[10] Previous research shows that part of this increase is due to the fact that the variance of the dependent variable increases at longer horizons.

5 Concluding remarks

In this paper we establish a surprisingly strong similarity in the empirical performance between pricing models based on long term consumption growth with time variation in aggregate consumption volatility and short-run models of idiosyncratic risk. The models not only deliver a similar fit on a given set of portfolios at different time horizons, they also generate almost the same expected returns and highly correlated mean squared errors. The match of two linear models is greatest at horizons of 8 quarters. The degree of correlation between predicted average biennial returns reaches then slightly more than 93 percent. While both models perform very similarly in long horizon time-series forecasts, they can not be considered as substitutes for each other. The long-term consumption risk specification fits with lower levels of estimated risk aversion parameter, however, it also implies a loss of degrees of freedom.

We argue that the documented similarity can be accounted for by the ability of the cross-sectional dispersion in consumption to explain the transitory component in aggregate consumption growth. Specifically, we find that consumption dispersion is a robust predictor of the transitory part of aggregate consumption growth rate with correlation of about 47.5 percent.

We explain these findings in a simple theoretical model in which aggregate uncertainty is a function of idiosyncratic uncertainty and only long-term consumption growth risk is priced. In this model, agents receive heterogeneous signals about their common long-term consumption prospects. An implication of this being that consumption dispersion is priced *empirically* not necessarily because markets are incomplete but because it helps identify the permanent component of consumption growth.

Our theoretical interpretation not only explains the link between pricing models based on long-term aggregate consumption growth and those with idiosyncratic risk, it also implies a strong cross-sectional correlation between assets' 'true' consumption growth betas and their dispersion betas: Assets with high exposure to ultimate consumption risk should deliver a low exposure to idiosyncratic consumption risk. We find strong evidence of this effect in the data.

6 References

[1] Bansal, Ravi, and Amir Yaron, 2004, Risks for the long run: A potential resolution of asset pricing puzzles, *Journal of Finance*, vol. LIX(4), pp. 1481-1509.

[2] Breeden, Douglas, 1979, An intertemporal asset pricing model with stochastic consumption and investment opportunities, *Journal of Financial Economics*, 7, pp. 265-296.

[3] Campbell, John Y., and Gregory N. Mankiw, 1989, International evidence on the persistence of economic fluctuations, *Journal of Monetary Economics*, vol. 54(3), pp. 591-621.

[4] Cochrane, John H., 2001, Asset pricing, *Princeton, NJ: Princeton University Press.*

[5] Constantinides, George, and Darrell Duffie, 1996, Asset pricing with heterogeneous consumers, *Journal of Political Economy*, vol. 104(2), pp. 219-240.

[6] Fama, Eugene F., and James D. MacBeth, 1973, Risk, return, and equilibrium: Empirical tests, *Journal of Political Economy*, vol. 81(3), pp. 607-636.

[7] Hansen, Lars Peter, 1982, Large sample properties of generalized method of moments estimators, *Econometrica,* vol. 50(4), pp. 1029-1054.

[8] Hansen, Lars Peter, and Ravi Jagannathan, 1997, Assessing specification errors in stochastic discount factor models, *Journal of Finance*, vol. LIX(2), pp. 557-590.

[9] Hansen, Lars Peter, and Kenneth J. Singleton, 1982, Generalized instrumental variables estimation of nonlinear rational expectations models, *Econometrica,* vol. 50(5), pp. 1269-1288.

[10] Heaton, John, and Deborah J. Lucas, 1996, Evaluating the effects of incomplete markets on risk sharing and asset prices, *Journal of Political Economy*, 104(3), pp. 443-487.

[11] Jacobs, Kris, and Kevin Q. Wang, 2004, Idiosyncratic consumption risk and the cross section of asset returns, *Journal of Finance*, vol. LIX(5), pp. 2211-2252.

[12] Jagannathan, Ravi, and Yong Wang, 2007, Lazy investors, discretionary consumption, and the cross-section of stock returns, *Journal of Finance*, vol. LXII(4), pp. 1623-1661.

[13] Kandel, Shmuel, and Robert F. Stambaugh, 1990, Expectations and volatility of consumption and asset returns, *The Review of Financial Studies,* vol. 3(2), pp. 207-232.

[14] Lucas, Robert, 1978, Asset prices in an exchange economy, *Econometrica,* vol. 46(6), pp. 1429-1445.

[15] Lustig, Hanno, and Adrien Verdelhan, 2007, The cross section of foreign currency risk premia and consumption growth, *The American Economic Review,* vol. 97(1), pp. 89-117.

[16] Parker, Jonathan A., 2001, The consumption risk of the stock market, *Brookings Papers Economic Activity,* vol. 2, pp. 279-333.

[17] Parker, Jonathan A., 2003, The consumption risk and expected stock returns, *NBER Working Paper* 9548.

[18] Parker, Jonathan A., and Christian Julliard, 2005, Consumption risk and the cross section of expected returns, *Journal of Political Economy*, vol. 113(1), pp. 185-221.

[19] Rubinstein, Mark, 1976, The valuation of uncertain income streams and the pricing of options, *Bell Journal of Economics and Management Science,* 7, pp. 407-425.

[20] Sarkissian, Sergej, 2003, Incomplete consumption risk sharing and currency risk premiums, *The Review of Financial Studies,* vol. 16(3), pp. 983-1005.

[21] Shanken, Jay, 1992, On the estimation of beta-pricing models, *The Review of Financial Studies,* vol. 5(1), pp. 1-33.

Table 1

SDF Estimation for the C-CAPM, LRC-CAPM and HC-CAPM at Different Return Horizons

	Panel A: C-CAPM				**Panel B: LRC-CAPM**				**Panel C: HC-CAPM**					
	GMM with prespecified weighting matrix		Efficient GMM		GMM with prespecified weighting matrix		Efficient GMM		GMM with prespecified weighting matrix			Efficient GMM		
Return horizon	$b_{\Delta c}$	[p-value]	$b_{\Delta c}$	J-Test [p-value]	$b_{\Delta c}$	[p-value]	$b_{\Delta c}$	J-Test [p-value]	$b_{\Delta c}$	b_{σ}	[p-value]	$b_{\Delta c}$	b_{σ}	J-Test [p-value]
h = 1	197.47	[0.00]	189.95	47.25	11.92	[0.00]	8.33	69.11	92.75	71.49	[0.00]	70.70	71.48	55.63
	(2.83)		(18.41)	[0.00]	(6.79)		(4.87)	[0.00]	(2.33)	(4.09)		(4.61)	(7.85)	[0.00]
h = 2	72.43	[0.00]	62.75	41.99	10.88	[0.00]	8.38	87.05	42.10	43.03	[0.00]	32.60	50.94	60.19
	(6.75)		(13.19)	[0.00]	(7.35)		(6.77)	[0.00]	(4.02)	(2.91)		(6.89)	(10.35)	[0.00]
h = 3	44.44	[0.00]	38.06	43.47	9.99	[0.00]	8.32	69.68	26.39	33.87	[0.00]	26.14	32.51	138.95
	(8.53)		(9.87)	[0.00]	(8.04)		(9.91)	[0.00]	(6.76)	(3.93)		(11.76)	(7.86)	[0.00]
h = 4	21.22	[0.00]	16.53	122.39	9.37	[0.00]	7.44	98.68	10.57	23.18	[0.00]	10.08	24.42	94.23
	(8.90)		(8.35)	[0.00]	(8.61)		(6.70)	[0.00]	(6.91)	(5.71)		(18.46)	(22.17)	[0.00]
h = 5	17.68	[0.00]	14.62	150.10	8.77	[0.00]	7.38	142.57	8.73	22.03	[0.00]	7.99	24.08	204.89
	(10.85)		(8.16)	[0.00]	(9.40)		(7.35)	[0.00]	(7.21)	(5.85)		(19.11)	(19.61)	[0.00]
h = 8	17.21	[0.00]	15.66	137.88	7.99	[0.00]	7.55	190.41	8.50	21.81	[0.00]	7.57	23.96	120.74
	(11.47)		(25.92)	[0.00]	(10.70)		(22.11)	[0.00]	(7.76)	(6.04)		(14.59)	(19.03)	[0.00]
h = 12	143.41	[0.00]	129.16	117.38	160.38	[0.00]	159.66	386.71	39.00	270.76	[0.00]	32.09	270.98	257.41
	(11.24)		(14.43)	[0.00]	(11.74)		(134.80)	[0.00]	(2.74)	(7.64)		(3.11)	(11.00)	[0.00]

Note: The table includes results from GMM estimation of the SDF of the C-CAPM, LRC-CAPM and the HC-CAPM. The pricing kernel representation is given by

$$m_{t+1}^{C} = 1 - b_{\Delta c}\Delta c_{t+1}, \; m_{t+1}^{LRC} = 1 - b_{\Delta c}\Delta c_{t+1}^{P}, \text{ and } m_{t+1}^{HC} = 1 - b_{\Delta c}\Delta c_{t+1} + b_{\sigma}\sigma_{K,t+1}^{2},$$

where Δc_{t+1} denotes short-run world consumption growth rate in the case of the C-CAPM and HC-CAPM, Δc_{t+1}^{P} denotes the permanent component of long-run world consumption growth rate in the case of the LRC-CAPM, and $\sigma_{K,t+1}^{2}$ denotes short-run cross-country consumption dispersion. Column *J*-Test gives the Hansen's goodness-of-fit χ^2-statistics. The matrix of instruments is composed of a constant and lagged world returns. The Newey-West corrected *t*-statistics are reported in parentheses below the estimates; *p*-values of the *J*-test for the overidentifying restrictions are provided in brackets. All returns are quarterly rates.

Table 2
Hansen-Jagannathan distance measure

Model	*HJ-distance:* $\min_{m \in M} \| m - m^x \|$
C-CAPM	1.4577
LRC-CAPM	0.9600
HC-CAPM	0.9148

Note: The table represents the results from GMM estimation of the C-CAPM, LRC-CAPM, and the HC-CAPM. The pricing kernel representation is given by

$$m_{t+1}^{C} = 1 - b_{\Delta c} \Delta c_{t+1}, \; m_{t+1}^{LRC} = 1 - b_{\Delta c} \Delta c_{t+1}^{P}, \text{ and } m_{t+1}^{HC} = 1 - b_{\Delta c} \Delta c_{t+1} + b_{\sigma} \sigma_{K,t+1}^{2},$$

where Δc_{t+1} denotes short-run world consumption growth rate in the case of the C-CAPM and HC-CAPM, Δc_{t+1}^{P} denotes the permanent component of long-run world consumption growth rate in the case of the LRC-CAPM, and $\sigma_{K,t+1}^{2}$ denotes short-run cross-country consumption dispersion. All returns are quarterly rates.

Table 3
HC-CAPM and LRC-CAPM

Return Horizon	γ_0	γ_1	$R^2(\%)$	$\overline{R}^2(\%)$
h = 1	0.0079 (0.8308)	0.8500*** (2.6965)	52.75	46.00
h = 2	0.0146 (0.9066)	0.8236*** (3.4661)	59.26	53.44
h = 3	0.0133 (0.7625)	0.9287*** (5.7720)	65.32	60.36
h = 4	0.0012 (0.0905)	0.9385*** (12.6392)	77.24	73.98
h = 5	-0.0258 (-1.5925)	1.1050*** (11.5030)	92.42	91.34
h = 8	-0.0135 (-0.8002)	1.0718*** (16.2077)	82.81	80.35
h = 12	0.0541 (0.9343)	1.4617*** (10.6920)	83.91	81.61

Note: The table represents the results from OLS regressions of mean returns predicted by the HC-CAPM on mean returns predicted by the one-factor LRC-CAPM: $\overline{R}_k^{HC} = \gamma_0 + \gamma_1 \overline{R}_k^{LRC} + \varepsilon_k$. In parentheses under the estimates are *t*-statistics. Significance at the 10% level is denoted by $.^{*}$, at the 5% level by $.^{**}$, and at the 1% level by $.^{***}$. The last two columns give the R^2 and the adjusted R^2. Both models are estimated by GMM (see Table 1).

Table 4
LRC-CAPM: First stage Fama-MacBeth

h	1	4	8	12
Panel A: Pooled OLS Regression Estimates				
β_h	1.0104***	0.9946***	0.8866***	0.7057***
	(109.44)	(36.074)	(17.094)	(7.961)
R^2	0.0132	0.0638	0.1306	0.1519
Panel B: Time Series OLS Regression Estimates				
β_h^{CAN}	0.268	1.4245*	3.1742***	104.1268***
	(0.7724)	(1.9586)	(4.9026)	(5.5684)
R^2	0.0107	0.0624	0.18	0.1953
β_h^{FR}	0.9034**	4.029***	8.6968***	275.0973***
	(2.4288)	(4.6207)	(5.55)	(6.3851)
R^2	0.0644	0.2194	0.3394	0.3248
β_h^{GER}	0.3107	1.7355**	4.1986***	122.666***
	(0.9263)	(2.0314)	(2.8803)	(3.1296)
R^2	0.011	0.0471	0.1146	0.1091
β_h^{ITL}	0.879*	4.0308***	8.8191***	306.5584***
	(1.7613)	(3.0396)	(3.3339)	(3.9878)
R^2	0.0398	0.1022	0.1477	0.1769
β_h^{JAP}	0.4424	2.3541**	7.1535***	371.1458***
	(1.0486)	(2.4686)	(4.0309)	(5.2454)
R^2	0.0154	0.082	0.201	0.3283
β_h^{SWITZ}	0.1232	0.915	2.486*	54.5038
	(0.3661)	(1.2444)	(1.8867)	(1.2972)
R^2	0.0018	0.0187	0.0545	0.0267
β_h^{UK}	0.3754	1.7649***	4.6291***	175.9115***
	(1.2344)	(3.1628)	(5.3741)	(5.673)
R^2	0.0165	0.1062	0.3272	0.3589
β_h^{US}	-0.0741	0.0134	0.289	9.369
	(-0.3189)	(0.0283)	(0.5144)	(0.554)
R^2	0.0013	0.01	0.0039	0.004

Note: The table gives estimates from first stage Fama and MacBeth (1973) pooled panel and single country time series regressions of the form:

$$r_{t+h}^{k} = \alpha_h + \beta_h \Delta c_{t+h}^{P} + \varepsilon_{t+h} \quad \text{(Panel A)}$$

and

$$r_{t+h}^{k} = \alpha_h^{k} + \beta_h^{k} \Delta c_{t+h}^{P} + \varepsilon_{t+h}^{k} \quad \text{(Panel B)},$$

where r_{t+h}^{k} is return of country *k* at time *t+h* and Δc_{t+h}^{P} denotes the permanent component of long-run world consumption growth rate. Excess returns at horizon *h* are obtained by summing up logarithmic one-period returns. In parentheses under the estimates are Valkanov-*t*-statistics corrected for small sample size. Adjusted R^2 is given below parameter estimates. Significance at the 10% level is denoted by $.^{*}$, at the 5% level by $.^{**}$, and at the 1% level by $.^{***}$.

Table 5
HC-CAPM: First stage Fama-MacBeth

h	1	4	8	12
Panel A: Pooled OLS Regression Estimates				
$\beta_{1,h}$	-87.329**	-54.726**	-159.56***	-134.10**
	(-2.159)	(-2.130)	(-3.941)	(-2.106)
$\beta_{2,h}$	1.706**	-1.811***	-5.2056***	-105.3***
	(2.171)	(-3.523)	(-6.483)	(-10.757)
R^2	0.0057	0.0759	0.1568	0.1925
Panel B: Time Series OLS Regression Estimates				
$\beta_{1,h}^{CAN}$	-0.0816	3.0588***	4.4623***	45.8945***
	(-0.0494)	(3.1627)	(3.7557)	(3.8841)
$\beta_{2,h}^{CAN}$	0.0051	0.0428	0.0895*	0.0737
	(0.4923)	(1.523)	(1.8193)	(1.2571)
R^2	0.0025	0.0958	0.1274	0.0881
$\beta_{1,h}^{FR}$	-0.9398	5.0103***	10.956***	165.9542***
	(-0.41)	(3.1439)	(4.7304)	(6.8486)
$\beta_{2,h}^{FR}$	-0.0147	-0.0635	-0.0978	0.0174***
	(-1.044)	(-1.4627)	(-1.1296)	(0.1107)
R^2	0.0099	0.1706	0.2581	0.316
$\beta_{1,h}^{GER}$	-0.559	2.0521*	6.8713***	122.8463***
	(-0.2577)	(1.7836)	(3.9552)	(6.4617)
$\beta_{2,h}^{GER}$	-0.0165	-0.1088***	-0.1917***	-0.0868
	(-1.4449)	(-2.7457)	(-3.1156)	(-0.7688)
R^2	0.0183	0.1158	0.2565	0.3624
$\beta_{1,h}^{ITL}$	-0.4469	3.1873	5.7754	119.43386**
	(-0.187)	(1.4701)	(1.5731)	(2.9162)
$\beta_{2,h}^{ITL}$	-0.007	-0.1421*	-0.36***	-0.4519***
	(-0.4167)	(-2.0489)	(-2.6156)	(-1.6352)
R^2	0.0016	0.0843	0.1127	0.137
$\beta_{1,h}^{JAP}$	1.2879	5.2266***	13.1788***	235.735***
	(0.6295)	(4.2752)	(6.0723)	(6.3036)
$\beta_{2,h}^{JAP}$	0.0067	-0.0616	-0.1515**	-0.1993
	(0.4668)	(-1.4498)	(-2.0951)	(-1.1037)
R^2	0.0045	0.2075	0.3537	0.4352
$\beta_{1,h}^{SWITZ}$	-1.4355	2.063*	5.3115***	75.3362***
	(-0.7622)	(1.8934)	(3.3397)	(3.7134)
$\beta_{2,h}^{SWITZ}$	-0.0207*	-0.1043***	-0.2242***	-0.3015***
	(-1.8378)	(-3.2744)	(-4.8797)	(-2.8481)
R^2	0.0292	0.1546	0.2889	0.2887
$\beta_{1,h}^{UK}$	1.2118	2.6324**	6.8596***	115.2801***
	(0.4639)	(1.9899)	(5.7062)	(6.6503)
$\beta_{2,h}^{UK}$	-0.0061	0.0011	-0.0242	0.075
	(-0.4692)	(0.0341)	(-0.5766)	(0.8485)
R^2	0.005	0.0567	0.2714	0.3131
$\beta_{1,h}^{US}$	-0.7046	0.3769	1.6953*	11.029
	(-0.4948)	(0.44)	(1.9297)	(1.1097)
$\beta_{2,h}^{US}$	0.0017	0.00	0.0016	-0.0158
	(0.1886)	(0.0012)	(0.0484)	(-0.2474)
R^2	0.0031	0.0029	0.0405	0.0168

Note: The table gives estimates from first stage Fama and MacBeth (1973) pooled panel and single country time series regressions of the form:

$$r_{t+h}^{k} = \alpha_h + \beta_{1,h}\Delta c_{t+h} + \beta_{2,h}\sigma_{K,t+h}^{2} + \varepsilon_{t+h} \quad \text{(Panel A)}$$

and
$$r_{t+h}^{k} = \alpha_h^k + \beta_{1,h}^k\Delta c_{t+h} + \beta_{2,h}^k\sigma_{K,t+h}^{2} + \varepsilon_{t+h}^k \quad \text{(Panel B),}$$

where r_{t+h}^{k} return of country *k* at time *t*+*h*, Δc_{t+h} denotes the short-run world consumption growth rate and $\sigma_{K,t+h}^{2}$ denotes likewise short-run consumption dispersion across *K* countries. Excess returns at horizon *h* are obtained by summing up logarithmic one-period returns. In parentheses under the estimates are Valkanov-*t*-statistics corrected for small sample size. Adjusted R^2 is given below parameter estimates. Significance at the 10% level is denoted by $.^{*}$, at the 5% level by $.^{**}$, and at the 1% level by $.^{***}$.

Table 6
Second stage Fama-MacBeth

	α_0	$\alpha_{\Delta c}$	α_σ	
Panel A: Results for the C-CAPM				
Coefficient	0.03123***	0.00339		
t (White-LS)	16.261	2.947		$R^2 = 12.63\%$
t (Shanken)	14.34	2.60		
Panel B: Results for the HC-CAPM				
Coefficient	0.0284***	0.00457**	-0.4815***	
t (White-LS)	11.807	3.090	-2.792	$R^2 = 54.43\%$
t (Shanken)	9.64	2.52	-2.28	
Panel C: Results for the LRC-CAPM				
Coefficient	0.0281***	0.00923**		
t (White-LS)	10.206	2.2336		$R^2 = 17.59\%$
t (Shanken)	9.00	1.97		
Panel D: Results for the Two-Factor C-CAPM: Δc^P_{t+1} and $\sigma^2_{K,t+1}$				
Coefficient	0.0256***	0.01058*	-0.3140	
t (White-LS)	12.590	3.4063	-1.5721	$R^2 = 21.73\%$
t (Shanken)	10.28	2.78	-1.10	

Note: The table represents the results from second stage Fama and MacBeth (1973) cross-sectional regressions. The *t*-statistics presented are: *t*(White-LS), from the White LS regression and *t*(Shanken), which adjusts for heteroskedasticity and the moving average process induced by the overlapping observations. Significance at the 10% level is denoted by .* , at the 5% level by .** , and at the 1% level by .*** .

Table 7
Forecastability of the Dispersion Factor

Differencing Horizon	β_s	*t*-statistic	$R^2(\%)$	$\overline{R}^2(\%)$
s = 2	-7.15	-1.580	4.880	3.80
s = 3	-10.388**	-2.504	7.15	6.26
s = 4	-10.858***	-2.975	8.48	7.42
s = 8	-16.299***	-3.594	13.29	12.23
s = 10	-16.859**	-3.5816	13.07	11.98
s = 12	-18.473***	-3.8004	14.24	13.14
s = 13	-18.282***	-3.6823	13.67	12.55

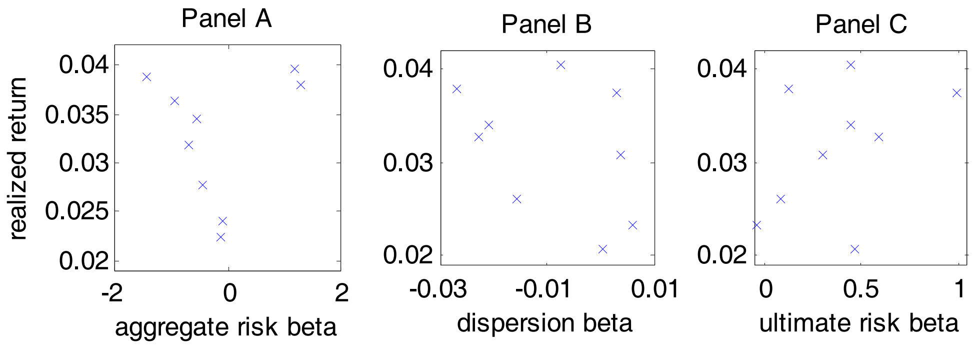
Note: The table represents the results of forecasting regression of transitory component in consumption growth on short-run consumption dispersion factor: $\Delta c_{t+1}^{T}(s) = \beta_0 + \beta_s \sigma_{K,t+1}^{2} + \varepsilon_{t+s}$. The transitory component of consumption growth is measured as a difference between aggregate and permanent components of *s*-quarterly long-run consumption growth rate and $\sigma_{K,t+1}^{2}$ denotes the quarterly cross-country consumption dispersion measured as cross-sectional variance of consumption growth rates. The permanent part of consumption growth is calculated as a quarterly equivalent of the ultimate GDP-weighted world per capita consumption growth over *s* periods. The *t*-statistic is the Newey-West adjusted. Significance at the 10% level is denoted by $.^{*}$, at the 5% level by $.^{**}$, and at the 1% level by $.^{***}$. The last two columns give the R^2 and the adjusted R^2.

Table 8
Partial Correlations of Transitory Consumption Growth and Dispersion Factor

Differencing Horizon	$\rho\left(\Delta c_{t+1}^{T}(s), \sigma_{K,t+1}^{2}\right)$	*p*-value
s = 2	-0.2209	0.0364
s = 3	-0.2707	0.0103
s = 4	-0.2912	0.0059
s = 8	-0.3645	0.0007
s = 10	-0.3615	0.0008
s = 12	-0.3774	0.0006
s = 13	-0.3697	0.0008

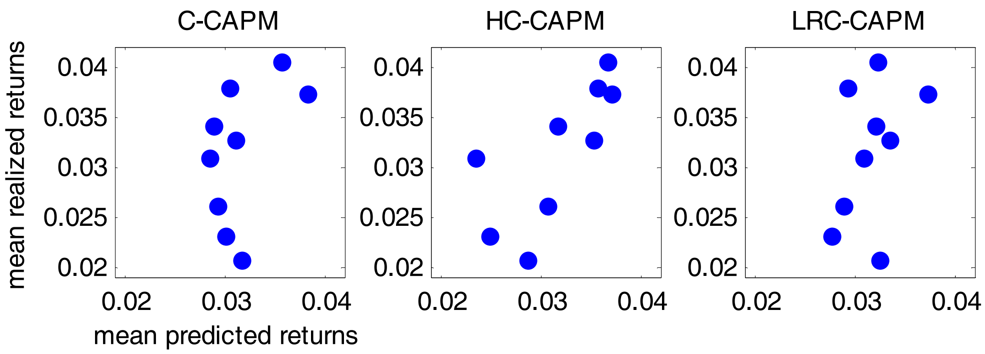
Note: The table represents the correlation coefficients of transitory component in consumption growth and short-run cross-country consumption dispersion factor. The column *p*-value tests the hypothesis of no correlation. The transitory component of consumption growth is measured as a difference between aggregate and permanent components of *s*-quarterly long-run consumption growth rate and $\sigma_{K,t+1}^{2}$ denotes the quarterly cross-country consumption dispersion measured as cross-sectional variance of consumption growth rates. The permanent part of consumption growth is calculated as a quarterly equivalent of the ultimate GDP-weighted world per capita consumption growth over *s* periods.

Figure 1



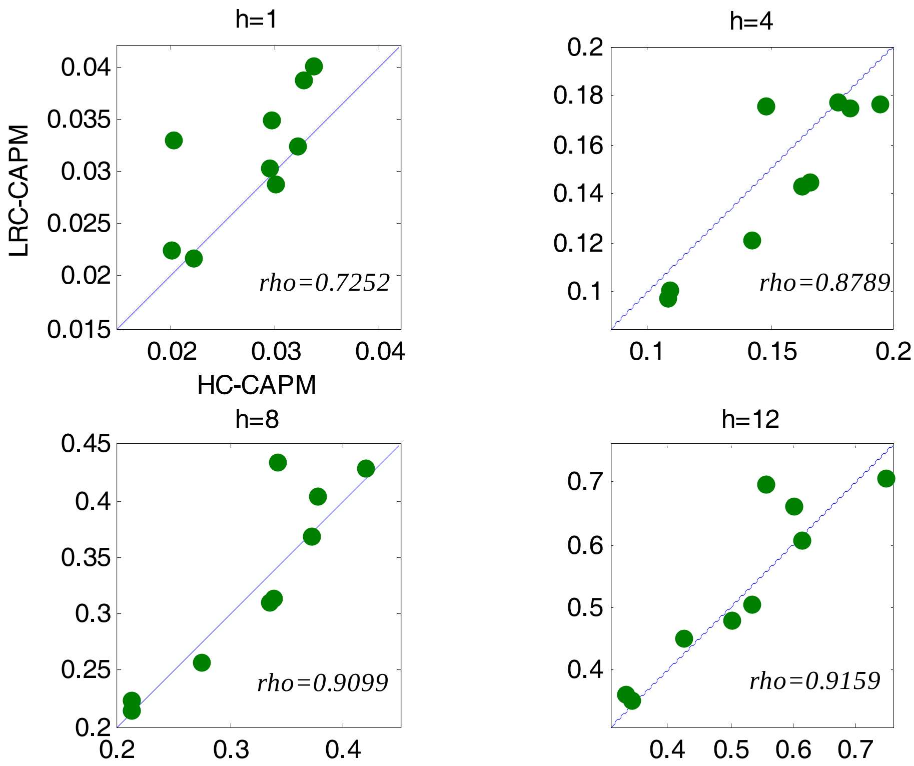
Note: Average realized returns (vertical axis) are plotted against estimated betas (horizontal axis) based on cross-sectional regressions for different consumption risk models. All returns are quarterly rates. The betas are estimated by Fama-MacBeth.

Figure 2



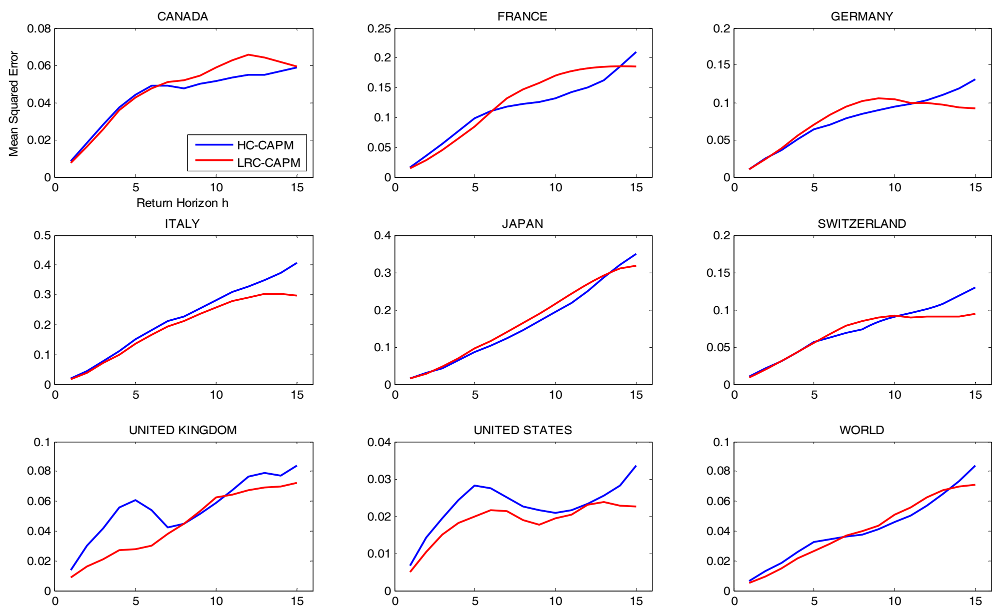
Note: Average realized returns (vertical axis) are plotted against average predicted returns (horizontal axis) based on cross-sectional regressions for different consumption risk models. All returns are quarterly rates. The returns are estimated by Fama-MacBeth.

Figure 3



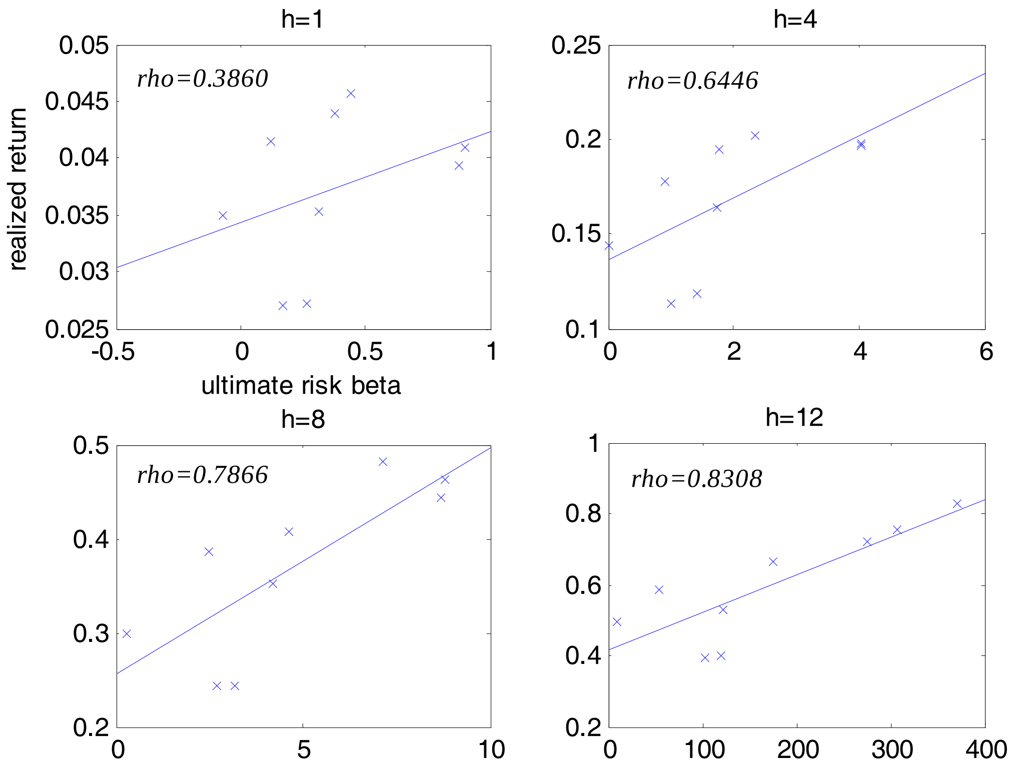
Note: This figure compares average fitted returns of the short-run two-factor HC-CAPM (horizontal axis) and average fitted returns of the one-factor LRC-CAPM (vertical axis) for different holding period returns (h = 1, 4, 8 and 12 quarters). The models are estimated by GMM with instrument vector composed of a constant and lag of 12 used in the weighting matrix (see Table 1).

Figure 4



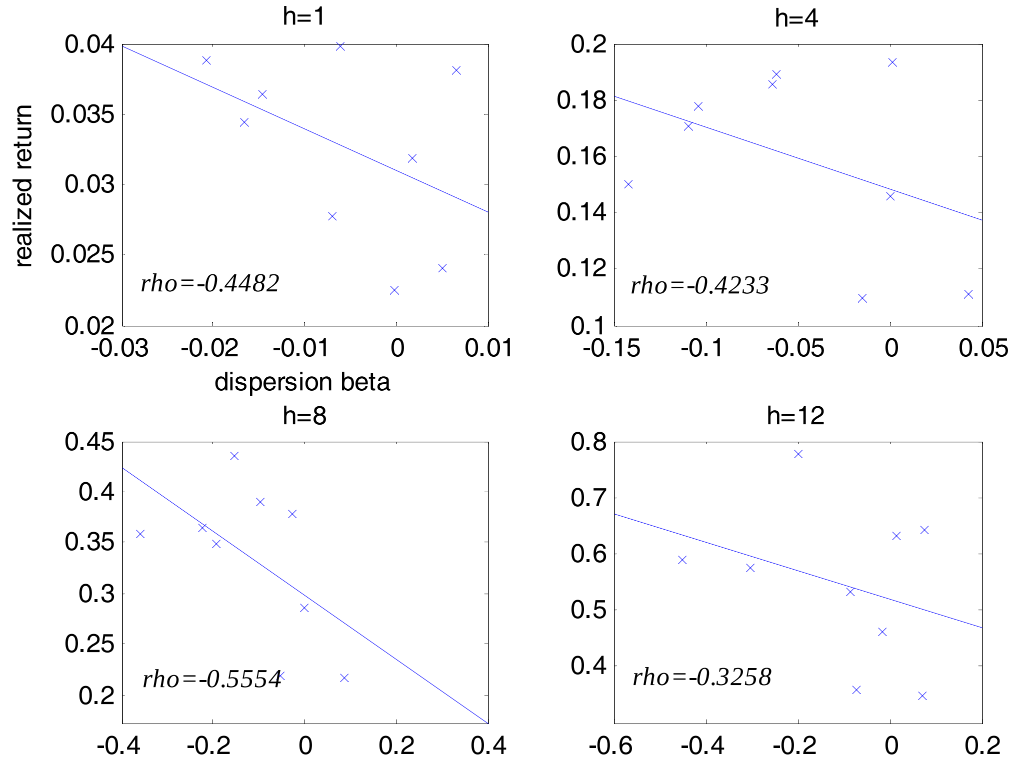
Note: This figure compares mean squared errors obtained from the short-run two-factor HC-CAPM (blue line) and one-factor LRC-CAPM (red line) for different holding period returns (h = 1, 2, …, 15 quarters).

Figure 5



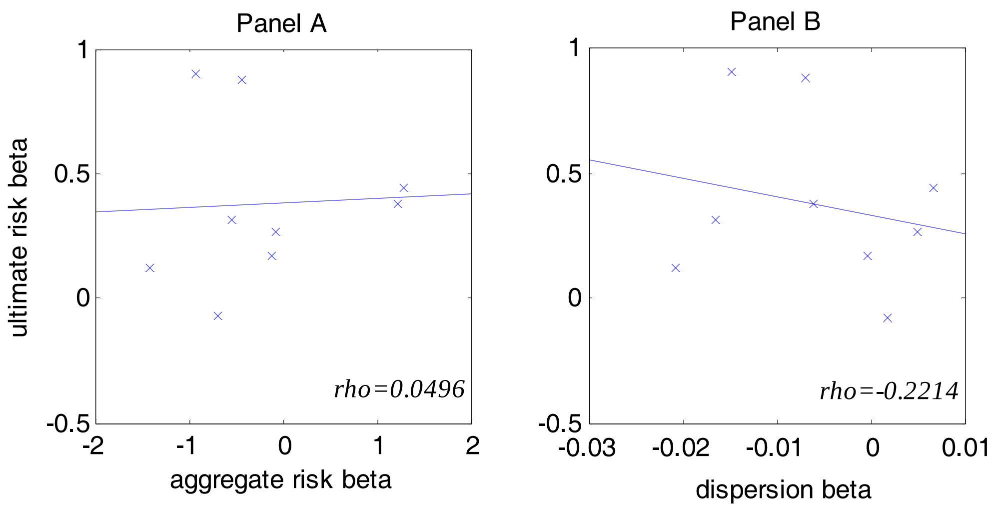
Note: The figure plots the estimated ultimate consumption risk betas (horizontal axis) from the one-factor LRC-CAPM against average realized returns (vertical axis). The model is estimated by Fama-MacBeth at different time horizons (h = 1, 4, 8, and 12 quarters). The estimated betas are those from Table 3.

Figure 6



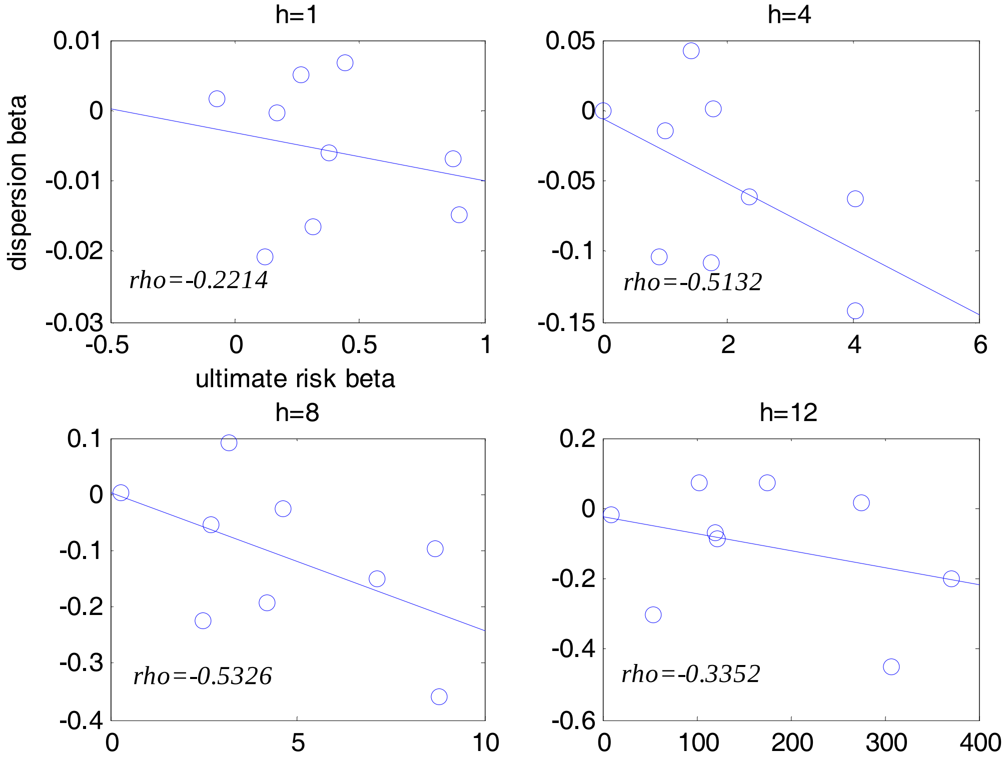
Note: The figure plots the estimated dispersion betas (horizontal axis) from the short-run two-factor HC-CAPM against average realized returns (vertical axis). The model is estimated by Fama-MacBeth at different time horizons (*h* = 1, 4, 8, and 12 quarters). The estimated betas are those from Table 4.

Figure 7



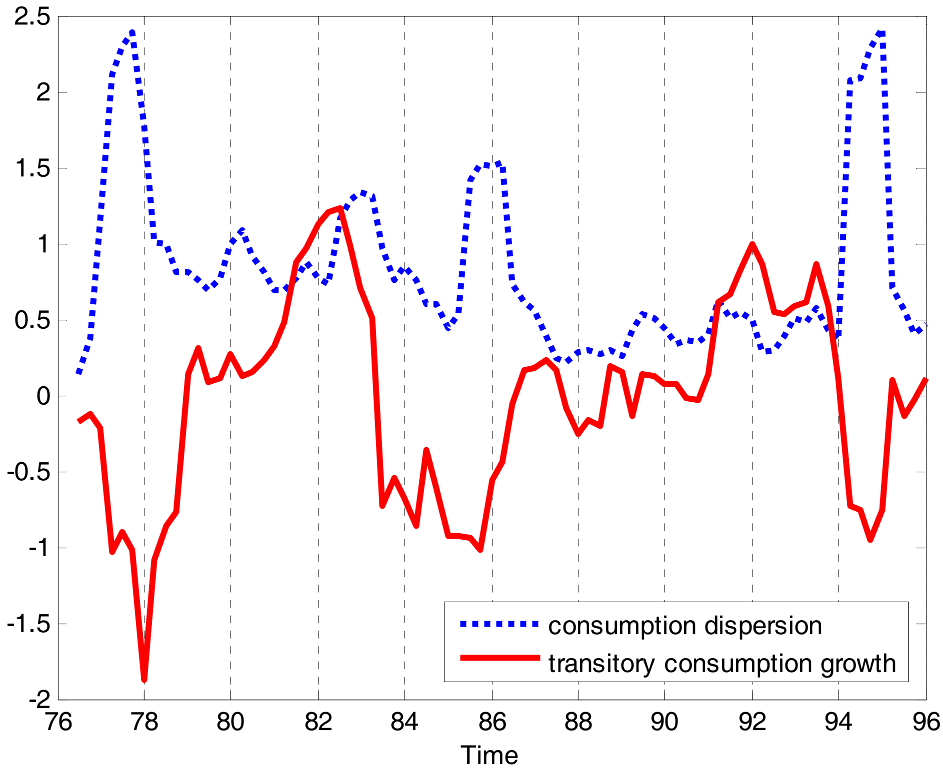
Note: Panels A and B plot estimated aggregate risk and dispersion betas from the short-run two-factor HC-CAPM (horizontal axis) against estimated ultimate risk beta (vertical axis) from the long-run one-factor LRC-CAPM. Ultimate risk betas are those from Table 3. Dispersion betas are those from Table 4.

Figure 8



Note: Estimated ultimate risk betas (horizontal axis) are plotted against estimated dispersion betas (vertical axis). Both models are estimated by Fama-MacBeth at different time horizons (*h* = 1, 4, 8, and 12 quarters). Beta estimates are those from Figures 5 and 6.

Figure 9



Notes: This figure plots the transitory consumption growth component (bold line) along with the short-run consumption dispersion factor (dotted line). Both series are standardized. Contemporaneous correlation of two series is -0.475.