Evidence for state and time non-separable preferences: The case of Finland[§]

Nader Shahzad Virk[*]

January 15, 2012

Abstract

Preferential modifications to the standard state and time separable power utility are studied for the Finnish equity and bond returns. The reported ambivalence of the high equity premium and low Sharpe ratio makes the Finnish market an important case study. The estimations of the Epstein and Zin (1991) recursive utility and the Campbell and Cochrane (1999) habit formation preferences show that Finnish risk premia are time-varying across samples. Moreover, the results demonstrate that stronger time preferences improve the explanation of asset returns for the modified preferences more so than assuming tighter time preference and higher risk aversion. We conclude that the Campbell-Cochrane-based pricing kernel outperforms the competing models in generating plausible model parameters and suppressing specification errors. The study supports the U.S. evidence relative to the conclusions drawn from the European economies in comparable studies.

JEL Classifications: C32, G12, G15.

Keywords: power utility, iterated GMM, risk premia, habit aggregation, risk factor, specification errors.

[§] The GMM code library from Mike Cliff in MATLAB is used in model estimations. This paper benefitted from comments and suggestions from the participants at Hanken School of Economics departmental seminars in 2010 and 2011 (Helsinki and Vaasa); GSF Winter research seminar in 2010 at Lappeenranta University; the Arne Ryde workshop in financial economics in 2011 at Lund University, Sweden; and the 18th annual MFS society conference in Rome, Italy. The comments and suggestions from Anders Löflund, Johan Knif, Peter Nyberg, Ihsan Badshah, and Jason Wei were especially helpful. Earlier versions of this paper have been presented under title "Canonical asset pricing models with risk separability and habit formation: Evidence from Finland."

[*] Contact author address: Hanken School of Economics, Department of Finance and Statistics, P.O. Box 479, 00101 Helsinki, Finland. Email: nadir.virk@hanken.fi.

1. Introduction

The equity premium puzzle of Mehra and Prescott (1985) questioned why the consumption-based general equilibrium model (CCAPM) failed to explain the larger return on stocks in comparison to the riskless asset in the economy, despite no particular differences in their correlations with systematic consumption risk. Henceforth, this question has been repeatedly addressed in the literature on the margins of macroeconomics and finance. The earliest empirical evidence (Hansen & Singleton, 1982) reported unfavorable results for the standard power utility-based CCAPM. The international evidence for the standard model (Cumby, 1990) yielded similar implications as the equity premium puzzle of Mehra and Prescott (1985), with the exception of the Japanese stock market (Hamori, 1992). The inability of the equilibrium models to explain the larger returns on volatile stocks brings into questions the viability of the models for quantitative assessment.

Nevertheless, the empirical inability of the model challenges the paradigms that characterize stocks as comparatively risky claims. Subsequently, the research has attempted to explain the equity premium in numerous ways. One way involves modifying preferences central to the standard model that restrict utility function to the constant relative risk aversion (CRRA) class.[1] The proposed adjustments include state non-separable preferential structure by Epstein and Zin (EZ hereafter, 1989, 1991) and Weil (1989).[2] The other notable modification allows for habit aggregation in the preferences – that is, the time non-separable utility function (e.g., Abel,1990; Constantinides, 1990; Campbell & Cochrane (CC hereafter), 1999). The modified preferences have greatly improved the empirical waning of CCAPM.[3]

Most studies incorporating alternative developments have focused on the U.S. market. Explanations of the equity premium puzzle from non-U.S. markets have been limited in comparison. Hyde and Sherif (2005) reported evidence supporting the power utility and habit formation models, among others, for the U.K. stock market. Chen and Ludvigson (2009)

[1] The standard CCAPM assumes the von Neumann Morgenstern utility, that is, separable in time and states.

[2] Agents smoothing consumption across various states of nature under the CRRA preferences also smooth consumption across periods; that is, they dislike growth. The coefficient of relative risk aversion (RRA) is the reciprocal of the elasticity of intertemporal substitution (EIS) under the standard CCAPM. The conventional time additive and state separable von Neumann-Morgenstern intertemporal utility function, under no a priori economic justification, relates investors' risk preferences with investors' consumption variations over time, specifically as reciprocals of one another.

[3] Otherwise, the model must conspicuously assume large relative risk aversion (RRA) values for the explanation of the observed equity premium. However, the assumption of large values for RRA has its own drawbacks, implying high subjective preferences under the model assumptions that cause interest rates to rise counterfactually – that is, risk the free rate puzzle of Weil (1999).

argued that within the consumption-based equilibrium models habit formation models are better candidates to describe the aggregate equity premium. Ghattassi (2008) concluded that in the long run, the CC (1999) consumption surplus ratio is a better predictor of asset returns than the proxies for consumption to wealth ratio (cay and cdy) by Lettau and Ludvigson (2001a, 2001b) for U.S. stocks.

Yet empirical evidence has surfaced recently for the alternative specifications coming from the developed European economies. Hyde et al. (2005) reported analytical evidence for the German and French stocks with the Hansen and Jagannathan (1991) non-parametric methodology. Engsted and Møller (2010) concluded that the CC (1999) habit formation model does not perform markedly better than the CCAPM for Danish stock and bond returns. Moreover, Engsted et al. (2010) showed that the surplus consumption ratio model does not improve on the standard CCAPM for numerous European economies in a post-World War II dataset. However, their evidence report log surplus consumption holds predictable information for future stock and bond returns. The reported evidence suggests an obvious divide between the results from the U.S. and the European economies.

1.1 Case Finland

Finland's stock market has been conventionally regarded as a developed market relying primarily on foreign trade. The market offers substantial annualized 8.8 percent (1990:01 to 2009:02) equity premium, a number usually associated with highly capitalized economies but with a comparably lower Sharpe ratio. The latter observation is often reported for European economies (see Hyde et al., 2005; Campbell, 1999) but with comparably lower excess market returns. The lower Sharpe ratio makes the relative risk aversion (RRA) small, which is consistent with the European evidence contrary to the larger values reported for the U.S. and U.K. stock markets (e.g., Mehra & Prescott, 2003; Hyde & Sheriff, 2005).

The significant value of the market capitalization to GDP ratio for the Finnish stock market, which is 57 percent (cf. Virk (2012a), is a reasonable proxy for the aggregate wealth to total consumption. The total private consumption in Finland accounts for almost half of the total GDP (Viitanen, 2004). Thus the Finnish market is suitable for testing the consumption-based equilibrium models, given the underlying assumptions. Viitanen (2004) and Virk (2012b) showed that the growth in Finnish consumption across different samples is negative. The latter study also highlighted the requirement of implausibly large RA values (71) to justify the observed equity premium. In a series of essays, Viitanen (2004) studied the general equilibrium models using Finnish consumption data. The estimated models in Viitanen's work are broad and capture internal and external habit effects along with the consumption durability parameter. These models provide a sharp contrast to the aforementioned alternative modifications that take only one habit behavior at a time and must overlook the durable consumption expenditures for maintaining separations between consumption and investment under the model regulatory conditions.

The intriguing artifacts of the Finnish market place it at the margin of capitalized and European economies. Furthermore, a clear divide in the empirical evidence from capital-intensive economies (such as the U.S. and the U.K.) and European economies calls for evidence from other markets. Therefore, this study approximates Finnish representative agents' temporal and state contingent risk preferences with EZ's (1991) recursive utility specification and the habit persistence specifications of Constantinides (1990), Abel (1990), and CC (1999). The estimation of habit formation models using Finnish data is the first empirical attempt and accordingly extends the Finnish asset pricing literature.

The specifications are estimated with quarterly data for the period from 1990:01 to 2009:02.[4] However, in comparison to other studies (Engsted & Møller, 2010; Engsted et al., 2010), o~~ur~~ work analyzes larger numbers of equilibrium models for the Finnish market. The noted studies have often compared only the performance of the CC (1999) surplus consumption ratio model with the benchmark CCAPM. Therefore, the comparison provided in this study more fully addresses developments in the related literature. The relative performance of the alternative specifications is tested with the Hansen and Jagannathan (HJ, 1997) distance metric.

The non-parametric calibration of Hansen and Jagannathan (1991) volatility bounds displays high levels of risk aversion that can be reduced relative to the standard model, consistent with the international evidence. Furthermore, the calibrations show that assuming higher subjective preference can also help explain related facts. The alternative utility specifications cannot be rejected with Hansen's (1982) over identification test. The EZ (1991) preferences show agents are more risk averse in the full sample than in the subsample estimations. The inferences from the CC (1999) estimations record similar movements in risk

[4] The sample period starts from 1990 instead of from 1975:01 because of the non-availability of aggregate dividend data that restricts the estimation of the Campbell and Cochrane (1999) external habit model. Furthermore, the model estimations are also reported for the sample, excluding the recession period – that is, from 1995:01 through 2009:02 – and are referred to as stable or recovery periods in the text.

premia as per EZ (1991). The Abel (1990) and Constantinides (1990) habit specifications estimated highly implausible risk and temporal parameters for the Finnish agent respectively.

The successful models can also predict returns on risk free rates with fewer errors than other modified habit specifications. The average imprecision in the projections is lower with the EZ (1991) preferences in the full sample, whereas CC (1999) outperforms all others in the subsample. Overall, the results from the EZ (1991) recursive utility provides mixed results that were sensitive to the selection of test assets, instruments, and sample periods. However, CC (1999) is the only model consistently producing economic implications for the subjective time and state preferences across all the specifications and samples. The better performance of the EZ (1991) and CC (1999) models produce lower HJ pricing errors. The EZ (1991) produces lowest distances for the specifications, including only stock portfolios, arguably by using returns on the aggregate stock index in the model stochastic discount factor (SDF). Otherwise, the CC (1999) habit model produces comparably smaller HJ pricing errors for all the remaining specifications.

The robustness tests with the linearized SDFs show that only the CC (1999) specification gets plausible risk signatures, which remain consistent even if we replace consumption growth with the dividend growth under the model assumptions of Lucas (1978). Importantly, the CC (1999) factor outperforms the market factor to capture variations in expected asset returns and command significant risk premia. We conclude that the CC (1999) external habit model lessens the severity of the puzzle for the Finnish market. The study reports evidence matching the studies for the U.S. (e.g., Chen & Ludvigson, 2009; Ghattassi, 2008) more so than the noted European evidence.

The remainder of the paper is organized as follows. The next section describes the data, and section 3 discusses alternate specifications used in the study to explain the asset return behavior and the equity puzzle. Sections 4 and 5 explain the estimations method and estimations output respectively. The last section concludes the study.

2. Model

We test the EZ (1989, 1991) state non-separable recursive preferences, internal habit model of Constantinides (1990) and external habit models of Abel (1990) and CC (1999). The functional formation of habit, which captures how current period consumption affects tomorrow's marginal utility of consumption, is defined arbitrarily in different studies. In the internal habit models, habit formation depends on agents' own past consumption patterns, whereas the external habit persistence evolves from aggregated consumption. The model-specific aggregated habit affects consumers' marginal utility exogenously in equilibrium. The functional form of the period utility function $U(C_t, X_t)$ is another modeling issue for the habit persistence specifications, where X_t, is the habit level. Abel specified utility as a power function of the ratio C_t/X_t , while Constantinides (1990) and CC (1999) stipulated utility as a power function of the difference – that is, $C_t - X_t$.

The specification of the habit in the period utility has an important effect on the behavior of the model risk aversion. The representative agent under the ratio preferential structure has constant RA, whereas risk aversion is time varying for the differenced utility specifications. The model moments are approximated under the assumption that all random processes are covariance stationary and their conditional and unconditional moments exist.

2.1. Epstein and Zin (1991) recursive utility

The tightly specified utility function of the standard CCAPM, under the additional constraint of the invertible relationship between RA and elasticity of intertemporal substitution (EIS), has been considered responsible for the dismal performance of the CCAPM. EZ (1989, 1991) and Weil (1989) developed state non-separable utility, which untied risk aversion from the EIS in consumption. EZ (1991) termed the modified recursive utility across the states as non-expected utility. The certain and random components of the utility specification are built on the models of Koopmans (1960) and Kreps and Porteus (1978) respectively. The recursive utility function is defined in infinite time settings such that

$$U_t = [(1-\beta)C_t^{\rho} + \beta(\widetilde{U}_{t+1}^{\alpha})^{\theta}]^{1/\rho}, \tag{1}$$

where $\beta = {}^{1}/_{1+\delta}$ is the discount factor given $\delta > 0$, and $\theta = \alpha/\rho$ where $1-\alpha$ is the coefficient RRA. EZ (1989) showed that the degree of risk aversion increases as α falls. The EIS, $\sigma = {}^{1}/_{1-\rho}$, along with rate of time preference (δ) is constant under the model assumptions providing deterministic future consumption. The earlier (late) resolution of uncertainty is preferred if $\alpha < (>)\rho$ or when θ < (>) 1. The power utility specification is nested in modified utility for $\theta = 1$.

The Euler equation is derived under the imposition of budget constraint (no labor income) that links consumption and wealth. This kind of budget constraint is valid under the complete market assumptions. The changes in aggregate stock market in the recursive preferences are proxy to shocks to wealth, which do not include return on human capital and therefore suffer from Roll's (1977) critique. The Euler representation implied by the EZ (1991) non-expected utility is

$$E_t\left[\{\beta(\tilde{C}_{t+1}/C_t)^{\rho-1}\}^{\theta}\{\tilde{R}_{w,t+1}\}^{\theta-1}(R_{i,t+1})\right]=1 \tag{2}$$

The intertemporal marginal rate of substitution (IMRS) implied by the EZ (1991) is a geometric weighted average of the SDF implied by the standard expected utility model and the SDF from the logarithmic expected utility model. The weights attached to each IMRS are determined through the ratio of risk parameter to the substitution parameter. The value $\theta = 1$ implies the consumption changes (the standard CCAPM) are enough to discount future payoffs on assets, whereas for $\theta = 0$ changes in the market index explains all – that is, the static CAPM case. For any other values of θ, both the nested specifications will contribute to determining expected asset returns. In order to maintain the identification of the parameter ρ, the following equation is also included in the model estimations, implying moment condition:

$$E_t[\{\beta(\tilde{C}_{t+1}/C_t)^{\rho-1}\tilde{R}_{w,t+1}-1\}^{\theta}/\theta]=0 \tag{3}$$

Equation (3) suggests an interesting case for $\rho \neq 0$ by distinguishing between the EZ implied logarithmic risk preferences and logarithmic expected utility model; for a detailed discussion, see EZ (1991: 268–270). The representative agent is assumed to have identical and homothetic preferences similar to the expected utility case.

2.2. Abel (1990) relative habit model

Abel (1990) studied the intertemporal decision problem for a representative agent with a generalized utility specification addressing both internal and external habit sensitivities. The model also nests the standard expected utility specification. The iso-elastic form for the period utility in Abel (1990) broke the time separability assumption and related consumption level with the prior period consumption levels (personal habit and the influence coming from others' consumption):

$$U_t=\sum_{j=0}^{\infty}\beta^j\frac{(C_{t+j}/X_{t+j-1})^{1-\alpha}}{1-\alpha}\quad,\alpha>0 \tag{4}$$

where X_{t+j} is a preference parameter such that $X_t = [c_{t-1}^{\rho} C_{t-1}^{1-\rho}]^{\theta}$, $\rho \geq 0$ and $\theta \geq 0$, while positing c_t as agents' own past consumption in period t and C_t as the aggregate per capita consumption in period t. The specified utility function converges to the standard expected utility formation for $\theta = 0$. Under the relative habit development, α measures the RRA for the representative agent and is constant.

Abel (1990) specified the external habit for $\rho = 0$ and $\theta = 1$. The constraints determine the degree of temporal non-separability with the aggregate per capita consumption at previous lag – that is, $X_t = C_{t-1}^{\theta}$. The utility specification is famously termed "catching up with the Joneses." The Euler equation in equilibrium under the specified utility is:

$$\beta E_t\left[(C_t/C_{t-1})^{\theta(\alpha-1)}(C_{t+1}/C_t)^{-\alpha}\, R_{i,t+1}\right] = 1 \tag{5}$$

The model-predicted risk free rate and risk premia are estimated assuming homoskedasticity and joint log normality (IID) of asset returns and consumption growth. Abel (1990) argued that relative consumption models could solve the large equity premium because the model implication allows for large risk aversion values without facing the Weil (1989) risk free rate puzzle. This fact is observed for positive values of time non-separability parameters but at the expense of counterfactually large fluctuations in the real risk free rate.

2.3. Constantinides (1990) internal habit model

Constantinides (1990) initiated the vast literature allowing for habit persistence. In the continuous time model, the internal habit/subsistence level is an exponentially weighted sum of past consumption levels. Moreover, consumption behavior for the representative agent is determined endogenously under the model assumptions in contrast to the standard CCAPM. The specification also nests the power utility specification (see equation (3) of Constantinides (1990)). The utility function is defined over the difference between current consumption and lagged past consumption such that

$$U_t = E_t\left[\sum_{j=0}^{\infty} \beta^j \frac{\left(C_{t+j}-X_{t+j}\right)^{1-\alpha}}{1-\alpha}\right] \tag{6}$$

where $X_{t+j} = \theta C_{t+j-1}$ with θ describing the persistence of habit into the utility function and higher values of θ manifest lower impact of the current consumption to influence investors' changing marginal utilities. The implied Euler equation by the continuous time model is:

$$E_t\left[\beta(\Delta c_{t-1})^{-\alpha}\frac{(\Delta c_t-\theta)^{-\alpha}-\beta\theta(\Delta c_t)^{-\alpha}E_{t+1}(\Delta c_{t+1}-\theta)^{-\alpha}}{(\Delta c_{t-1}-\theta)^{-\alpha}-\beta\theta(\Delta c_{t-1})^{-\alpha}E_t(\Delta c_t-\theta)^{-\alpha}}R_{i,t+1}\right]=1 \tag{7}$$

where $\Delta c_{t+i} = C_{t+i}/C_{t+i-1}$ for $i = -1,0,1$. The α is the utility curvature parameter such that $RRA = \frac{\alpha}{1-x/c}$, where x is the fixed subsistence level. The estimate for α will approximate the risk aversion when the ratio between consumption level and specified habit level nears one while always remaining below one. The model cannot resolve the equity premium puzzle without invoking large aversions to consumption risk. The model-implied Euler equation is with conditional operators; we assume consumption growth is IID to obtain the unconditional closed form solution in discrete time following Cochrane and Hansen (1992) and Hyde et al. (2005). Moreover, the predictions for the risk free rate are made assuming all the processes are jointly log normal. The derivation of the model-predicted return on risk free security is provided in Appendix A.

2.4. Campbell and Cochrane (1999) surplus consumption ratio model

CC (1999) proposed an external habit formation model devoid of the issues confronted with the conventional habit models. The agents in the model derive utility from a power function of the difference between consumption and a time-varying habit level such that

$$U(C_t, X_t) = \frac{(C_t-X_t)^{1-\alpha}}{1-\alpha}, C_t > X_t \tag{8}$$

where X_t is habit level depending on previous periods' consumption, and α is the utility curvature coefficient. The specified utility function allows time-varying, counter cyclical risk aversion equaling $\frac{\alpha}{S_t}$. Instead of modeling habit in the model, CC (1999) stipulated a process for the surplus consumption ratio: $S_t \equiv \frac{C_t-X_t}{C_t}$. The surplus consumption ratio is the fraction of consumption that exceeds habit in the prior periods and influences agents' marginal utility in making intertemporal decisions. The RRA rises in the model as consumption falls towards habit. The Euler equation under the first order condition is:

$$E_t\left[\beta\left(\frac{S_{t+1}}{S_t}\frac{C_{t+1}}{C_t}\right)^{-\alpha}R_{i,t+1}\right] = 1 \tag{9}.$$

The growth in the consumption levels is assumed to follow random walk process for the replication of the observed empirical facts regarding equity premium, predictability of asset returns, and stable and low levels of interest rates. The log surplus consumption following an AR (1) process – that is, $s_t \equiv \log(S_t)$ – is specified:

$$s_{t+1} = (1-\varphi)\bar{s} + \varphi s_t + \lambda(s_t)\nu_{c,t+1} \tag{10}$$

where $\Delta c_{t+1} = g + \nu_{c,t+1}$, such that $\nu_{c,t+1}$ is NIID consumption shocks with zero mean and constant variance. The parameter φ governs the persistence of the log surplus consumption ratio, and $\bar{s}$ is the steady state level of s_t. The sensitivity function $\lambda(s_t)$ controls the variations in the riskless interest rate for capturing the habit responses to consumption shocks with time-varying risk aversion and is specified as follows:

$$\lambda(s_t) = \begin{Bmatrix} \frac{1}{\bar{S}}\sqrt{1-2(s_t-\bar{s})} & if\ s_t \leq s_{max} \\ 0 & else \end{Bmatrix} \tag{11}$$

where $\bar{S} = \sqrt{\frac{\alpha\sigma_v^2}{1-\varphi}}$, $s_{max} \equiv \bar{s} + \frac{1}{2}(1-\bar{S}^2)$ and $\bar{s} = log(\bar{S})$.

The specification balances out the effects of "desire to borrow against future" and "precautionary savings." The power coefficient α controls the relationship between consumption growth and interest rates, while the risk aversion coefficient controls the price of risk. The model assumptions allow for high-risk aversion with low aversion to intertemporal substitution. The process for the surplus consumption ratio is unobserved and is central to the specification of the model. Therefore, the process in (10) is calibrated with the OLS estimates for g, σ_v^2, whereas, φ is the first order coefficient for the log price to dividend ratio pd. A grid search is performed for the parameter α, such that the observed return on riskless asset is estimated around the steady state surplus consumption ratio. The estimated steady state surplus consumption is taken as the initial value for calibrating the log surplus consumption ratio – that is, s at $t = 0$. The process is then estimated recursively across the samples such that the surplus consumption ratio is $S = exp(s)$.

3. Estimation method and the specification test

Let $\tilde{\boldsymbol{y}}$ be a SDF and $\boldsymbol{R}$ be the gross returns on *N* test assets. If the SDF correctly prices the *N* assets, the pricing errors $\boldsymbol{g}$ should satisfy the identity such that

$$g \equiv E_t[\tilde{y}_{t+1}R_{t+1} - 1_N] = 0_N \tag{12}.$$

The specification of $\tilde{y}$ is model dependent, and we only observe its empirical approximation in $\boldsymbol{y}$. In the context of this study, $\boldsymbol{y}$ corresponds to (2), (5), (7), and (9) for the CCAPM alternative specifications. The success of the empirical IMRS lies in the ability to approximate the identity in (12) as closely as possible and is dependent upon the unobservable model parameter vector $\boldsymbol{\theta}$. Let the model IMRS be a function of unobserved parameter vector and $\boldsymbol{Z_t}$ be a vector of K instruments observed in period t. Then the asset pricing model implies the following orthogonality conditions:

$$\boldsymbol{E_t}\left[\left(\boldsymbol{R_{t+1}y'_{t+1}(\theta) - 1_N}\right)\otimes\boldsymbol{Z_t}\right] = \boldsymbol{0_{NK}} \ . \qquad (13).$$

GMM estimates the pricing errors, with no information on the exact distribution of the model errors, from the sample analogous to the NK orthogonality conditions such that $\boldsymbol{g_T(\theta)} = \frac{\boldsymbol{1}}{\boldsymbol{T}}\sum_{\boldsymbol{t=1}}^{\boldsymbol{T}}\left[\left(\boldsymbol{R_{t+1}y'_{t+1}(\theta) - 1_N}\right)\otimes\boldsymbol{Z_t}\right] = \boldsymbol{0_{NK}}$ based upon T sample observations. The standard Hansen (1982) method estimates parameter vector in two steps. However, the approach for estimating parameter vector and then updating the weighting matrix can be iterated n times until the coefficients converge or the change in the objective function in (12) is sufficiently small. The estimations in the study are carried with the iterated GMM method. The parameter vector is estimated while minimizing the quadratic form of the type

$$J_T = g_T(\theta)' W g_T(\theta) \qquad (14)$$

where W is $NK \times NK$ weighting matrix and is model-dependent.[5] Nevertheless, the weighting matrix may also be model independent for the purpose of comparing different models, such as the identity matrix.[6]

The selection of the Hansen (1982) optimal weighting matrix is preferred in this study as it more often converges to plausible parameter solutions. Moreover, the weighting matrix also simplifies the hypothesis testing – that is, TJ_T converges to χ^2_{NK-q}, where q is the number of

[5] The spectral density matrix W=S $=\sum_{-\infty}^{\infty} g_T(\theta) g_{T-j}(\theta)'$ is the Newey and West (1987) hetroskedasticity and autocorrelation consistent (HAC) estimator. Bartlett kernel is used to downgrade the auto-covariance structure for the positive semi-definiteness of the S. The bandwidth algorithm is set to the Newey-West (1994) nonparametric method based on a truncated weighted sum of the estimated cross-moments, which control the number of auto-covariances in the HAC estimator and are important for consistent finite sample properties of S. The statistically optimal (most efficient) weighting matrix is obtained as the inverse of the covariance matrix of the sample orthogonality conditions – that is, S^{-1}. It provides the smallest possible standard errors whereas any suboptimal matrix may produce inconsistent estimates.

[6] The data/model independent weighting matrix may be kept constant across all the competing models, which is impossible with Hansen's (1982) optimal matrix. Otherwise, the selection of the identity matrix assigns equal weights to specified moment conditions and provides a basis for model comparison among different models. Engsted et al. (2010) used the identity matrix, but Garcia et al.'s (2005) study chose the optimal weighing to obtain the parameters with the smallest standard errors.

unknown parameters. Nevertheless, the value of quadratic minimization cannot be used to compare the relative size of pricing errors associated with different asset pricing models. In order to compare the performance of different models, we additionally compute Hansen and Jagannathan (1997) misspecification measures. The specification measure fixes the weighting matrix, in the GMM sense of equation (15), such that $\boldsymbol{W} = [\boldsymbol{E}(\boldsymbol{R}_{t+1}\boldsymbol{R}'_{t+1})]^{-1}]$. Hansen and Jagannathan (1997) showed that the type of quadratic form, as in (15), measures squared distance between candidate SDF $\boldsymbol{y}$ of a given model and the set of SDFs in admissible region that prices the N asset correctly. The square root of the squared distance is referred to as HJ-distance:

$$HJ - distance = \left(\left[E\left(R_{t+1}y_{t+1}(\theta)' - 1\right)' E(R_{t+1}{}'R_{t+1})'^{-1} E(R_{t+1}y_{t+1}(\theta)' - 1)\right]\right)^{1/2} (15).$$

The specification measure has an appealing economic interpretation such that HJ-distance=0.10 implies that the model-predicted asset prices deviate by 10 percent from the observed prices. Because the weighting matrix used in the estimations in (14) is suboptimal, the minimized value of the quadratic function does not converge to $\chi 2$ distribution. The asymptotic standard error for the HJ-distance is computed with delta method as described in Hansen et al. (1995):

$$v_t \equiv (y_{t+1})^2 - \ (y_{t+1} - \hat{\lambda}R_{t+1})^2 - 2E\iota'\hat{\lambda} \tag{16}$$

where y_{t+1} is the model specific IMRS and $\hat{\lambda}$ is the sample estimate of

$$\lambda = [E(R_{t+1}R'_{t+1})^{-1} \ (E(y_{t+1}R_{t+1}) - E\iota')]^{1/2} \tag{17}$$

The asymptotic distribution of $\widehat{HJ}$ is degenerative for testing the null hypothesis HJ=0. Thus, the asymptotic standard error gives a measure of precision of the $\widehat{HJ}$.

4. Data

The quarterly data extending from 1990:01 to 2009:02 is used for the model estimations. The series for nondurables and services (NDS) is provided by ETLA and is used as a proxy for aggregate private consumption. The aggregate consumption is divided by the total population to obtain the per capita consumption expenditure. The study uses five test assets, namely

(EURIBOR), 10-year government bonds (LGB), and Fama and French (1993) risk mimicking factors, Small-minus-Big (SMB) and High-minus-Low (HML), for analyzing the modified utility specifications. All the nominal series are deflated with implicit price deflator retrieved from the Statistics Finland database.

Table 1 provides summary statistics for the test variables and instruments used in the study. The negative consumption growth in the studied sample for the Finnish economy is a more distinctive feature then the positive growth across the markets (see Campbell, 1999).[7] The recession in the early 1990s yielded drastic drops in aggregate consumption, making the full sample consumption growth negative for the Finnish agent. The exclusion of the crisis period shows substantial recovery with a positive growth in the consumption growth. Therefore, estimations in the full period and the stable period provide a natural opportunity to explore the effects of fall and recovery in the consumption patterns in influencing Finnish representative agents' intertemporal choices.

The gross real returns (minus one) on stocks are much higher than the bond returns. The greater volatility of the stock returns and NDS implies larger confidence intervals for the corresponding averages than the bond returns, which tend to have significant mean returns. The annualized excess equity premium $4 \times (3.1\% - 0.009\%) = 8.8\%$ is substantial. The large Finnish equity premia are more of a capitalized economy feature, as reported for the U.S. and the U.K. markets, than the European economies. Hyde et al. (2005) reported annualized equity premiums of 2.28 percent and 1.79 percent for France and Germany, whereas Engsted and Møller (2010) reported an annualized premium of 4.33 percent for the Danish market. The sample Sharpe ratio of 11.3 percent is closer to the reported values of 11.3 percent and 10.1 percent for Germany and France respectively in Hyde et al. (2005) than to the U.S. ratio of 50 percent (Mehra & Prescott, 2003) and the U.K. ratio of 39.5 percent (Engsted, 1998).

The consumption growth for the sample period has more density on the left side causing the (negative) asymmetry, whereas other variables are generally positively skewed and leptokurtic (fat tailed). The normality hypothesis cannot be rejected for LGB and HML; otherwise, all remaining variables have excess kurtosis. Only the bond returns show signs of

[7] Similar observations are also noted for Finnish consumption growth in Oikarinen and Kahra (2002), Viitanen (2004), and Virk (2012b).

Table 1 Summary Statistics

The descriptive stats for the variables used in the study are reported such that μ and σ signifies quarterly average and standard deviation. All the averages are one minus the gross real return and the reported standard error (se) and 95% confidence interval are computed with $\sigma/\sqrt{T}$ and $\mu \pm 2 \times se$ respectively. The standard errors for skewness and kurtosis shown in parentheses are $(6/T)^{1/2}$ and $(24/T)^{1/2}$ respectively. Jarque-Bera (JB) test statistic, for testing under the null hypothesis whether the series is normally distributed, is distributed as χ^2 with 2 degrees of freedom. The significance for the autocorrelation coefficients are computed with Ljung Box (1978) Q-statistics at lag n, testing whether a group of n autocorrelations are significantly different from zero (under null hypothesis there is no autocorrelation in the residuals) or otherwise could also be interpreted as if the random variable is a white noise process. Q(n) is distributed as χ^2 with n degrees of freedom. The p-values significantly rejecting the null at 5% level and below are shown in bold.

	μ	se	[Confidence Interval]		Max.	Min.	σ	Skewness	Kurtosis	JB test	Prob.	*ρ(1)*	*ρ(2)*	*ρ(4)*	*ρ(6)*
R_f	**0.009**	0.001	[0.007	0.011	0.03	-0.004	0.009	0.91 (0.28)	3.29 (0.56)	11.02	0.00	**0.74**	**0.63**	**0.72**	**0.48**
LGB	**0.011**	0.001	[0.009	0.013]	0.03	-0.004	0.007	0.31 (0.28)	2.66 (0.56)	1.60	0.45	**0.64**	**0.52**	**0.69**	**0.47**
R_m	0.031	0.022	[-0.013	0.076]	0.90	-0.37	0.194	1.10 (0.28)	6.71 (0.56)	60.49	0.00	0.09	0.05	0.15	-0.01
SMB	0.013	0.013	[-0.012	0.037]	0.26	-0.16	0.112	1.71 (0.28)	8.52 (0.56)	137.49	0.00	-0.12	-0.00	0.11	0.12
HML	**0.001**	0.013	[-0.025	0.027]	0.29	-0.20	0.114	0.39 (0.28)	3.76 (0.56)	3.87	0.14	**0.27**	**0.10**	0.02	0.17
ΔC	-0.001	0.002	[-0.005	0.002]	0.03	-0.07	0.014	-1.16 (0.28)	7.17 (0.56)	74.06	0.00	-0.15	0.04	0.19	0.06
DY	**0.007**	0.0004	[0.0003	0.002]	0.07	0.01	0.01	0.66 (0.28)	4.74 (0.56)	15.32	0.00	**0.79**	**0.62**	**0.36**	**0.21**
PER	**0.04**	0.003	[0.033	0.043]	0.54	0.05	0.09	1.79 (0.28)	7.14 (0.56)	96.06	0.00	**0.80**	**0.60**	**0.43**	**0.24**

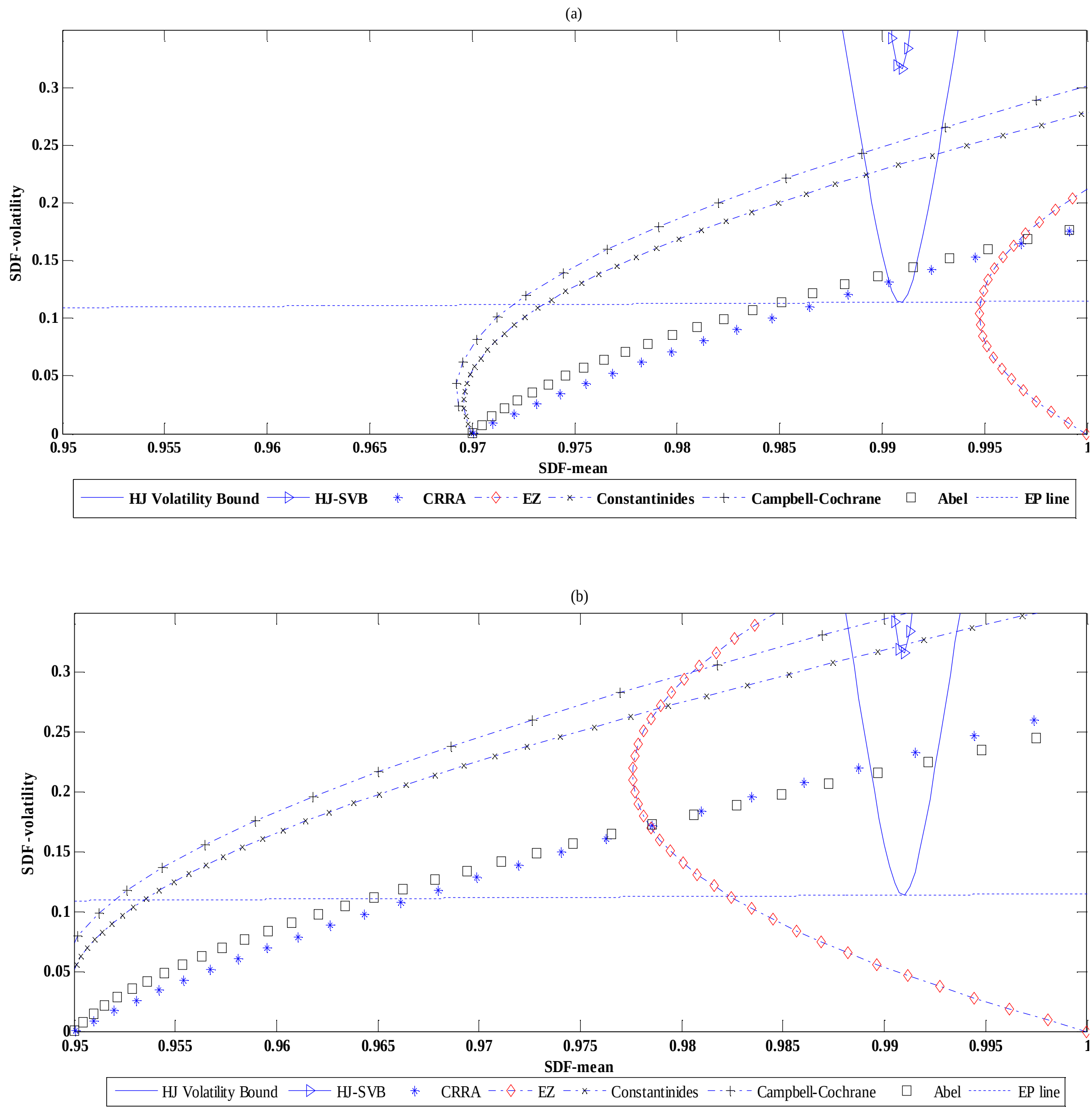

Figure1. The Hansen and Jagannathan (1991) minimum volatility bounds for Finnish asset data are stretched The pairs of SDF mean and volatility are generated for CCAPM and the preferential modifications developed on it while breaking the state and time separable assumptions under range of non-parametric solutions for model intertemporal substitution and risk aversion.

predictability. The study uses a number of instruments for estimating the model moment conditions. The instrument vectors are selected for which the Hansen (1982) moment identification test cannot be rejected. The selection of lagged aggregate price-to-earnings ratio and dividend yield is motivated for its higher predictability power in the asset pricing literature (e.g., Campbell & Shiller, 1988). Both instruments are rightly skewed and fat tailed, with high persistence in their levels. The selected instrument vectors are $(1\ R_{m,t}\ R_{m,t-1}\ R_{f,t}\ R_{f,t-1})$, $(1\ PER_t\ )$, and $(1\ PER_t\ DY_t)$, and are notated as INTI, INTII, and INTIII respectively across the model estimations.

5. Estimations

5.1. Hansen and Jagannathan (1991) min. volatility bounds

The nonparametric approach of Hansen and Jagannathan (1991) established a lower volatility bound on the model SDF. For the textbook treatment of the subject matter, we refer to Campbell, Lo, and Mackinlay (1997). The feasible regions for the mean and standard deviation of the employed models SDFs are the necessary precursors to the parametric estimations of the studied modifications to investors' preferential structure. The volatility bounds provide the analytical solutions for the model risk aversion under which the model SDFs possess the essential characterization to be consistent with asset return data.

The equity premium puzzle of Mehra and Prescott (1985) could also be reciprocated into the lower volatility bound of Hansen and Jagannathan (1991) for the model SDF.[8] The Hansen and Jangannathan (1991) volatility bounds are stretched, assuming represen~~ta~~tive agents' subjective preferences of 0.97 and 0.95, as shown in Figure 1 (a) and (b) respectively. The ρ parameter is 3 for EZ (1991) preferences, whereas the habit persistence 'θ' parameter values for the Abel (1990) and Constantinides (1990) model calibrations are assumed to be 1 and 0.5 respectively. The surplus consumption ratio is generated to reproduce the variations in the Finnish asset return data.

The locus (solid line) of the mean-volatility bound is drawn assuming perfect correlation between consumption growth and the asset returns (both stocks and bonds). The squeezed in volatility bound (solid line with arrows) manifests more realistic variability requirements in the model SDFs while accounting for the historical correlations. The theoretically motivated loci are quite liberal considering that the estimated projection for the correlation between the model SDF (standard CCAPM) and the aggregate index is only 17 percent, following the method stipulated in Hansen and Cochrane (1992). Therefore, the empirical SDFs must be

[8] The standard CCAPM model specifies the unconditional equity premium as $\boldsymbol{E}\left[\boldsymbol{R_i} - \boldsymbol{R_f}\right] = \boldsymbol{\alpha\sigma_{ic}}$, where α is the price of risk (risk aversion), and σ_{ic} implies the quantity of risk. Assuming the perfect correlation between consumption growth and aggregate wealth above, equality can also be represented as $\sigma_m \approx \alpha\sigma_c$. The σ_m is the model SDF volatility bounded from below to the maximum Sharpe ratio given asset return data (that is, 11.3 percent for the Finnish aggregate market), and σ_c is the consumption volatility. In summary, the model SDF should either be highly correlated with asset return data or, as shown in Hansen and Jagannathan (1991), should be highly volatile to cut through the minimum volatility bounds. Therefore, the smoothness of consumption growth implies large risk aversion to explain the growth and variability in the aggregate market index. Alternatively, for economically plausible risk aversion values, the consumption changes should exhibit larger variability. First, we do not observe any substantial volatility in the consumption changes. Second, if true, they should also induce large fluctuations in risk free rate (e.g., Weil, 1989), which has not been the case historically.

Table 2 Estimations with Epstein and Zin state non-separable preferences

Table reports the EZ (1991) state separable preferences such that first panel reports the estimation output for full sample and the 2nd panel reports for sub sample period. Specification I does the iterated GMM estimation using returns on risk free bond and long term bond and specification II tests only with stocks such that SMB and HML as test assets. The specification III takes the bond returns and stock returns together in one cross section for model testing. In each panel against the subsequent instrument vector model parameters for time preference, intertemporal substitution and ratio of risk parameter to the substitution parameter are reported with notations β, ρ, and θ. The standard errors for the model estimates are given in ().χ2-test is the Hansen's moment identification statistic reported with p-value in [].The predicted return on risk free rate with estimated parameters is reported against row r_f. Wald test tests if $H_o: \theta = 1$ to draw on the fact if EZ state separable preference fits the data better whereas, the non rejection of the implies from N-1 equations the standard power utility is still a reasonable approximation to describe agent behaviour.

	Panel I Full Sample			Panel II Sub Sample		
Specification I	INTI	INTII	INTIII	INTI	INTII	INTIII
β	**0.98**	**0.99**	**0.98**	**0.98**	**0.98**	**0.98**
	(0.002)	(0.002)	(0.001)	(0.007)	(0.005)	(0.004)
ρ	**1.37**	0.85	0.96	**-10.25**	**-6.92**	**-7.62**
	(0.28)	(1.13)	(0.52)	(3.66)	(1.64)	(1.85)
θ	**0.80**	**0.86**	**0.80**	**1.27**	**1.35**	**1.35**
	(0.02)	(0.06)	(0.02)	(0.28)	(0.41)	(0.33)
RRA	-0.10	0.26	0.23	14.02	10.35	11.27
χ2-test	13.20	7.14	9.94	10.03	8.97	9.38
p-value	[0.35]	[0.31]	[0.36]	[0.61]	[0.18]	[0.40]
r_f	0.014	0.012	0.014	0.003	0.01	0.008
Wald test	**92.29**	**5.81**	**81.64**	0.93	0.75	1.12
p-value	[0.00]	[0.02]	[0.00]	[0.34]	[0.39]	[0.29]
Specification II						
β	**0.99**	**0.98**	**0.98**	**0.98**	**0.99**	**0.99**
	(0.004)	(0.006)	(0.005)	(0.010)	(0.020)	(0.013)
ρ	0.97	-0.13	-1.47	**-12.08**	-12.87	**-9.09**
	(0.90)	(2.42)	(2.26)	(4.22)	(7.15)	(4.00)
θ	**0.77**	**0.60**	**0.58**	**1.70**	0.36	0.29
	(0.08)	(0.17)	(0.11)	(0.29)	(1.35)	(0.86)
RRA	0.26	1.08	1.86	21.54	5.62	3.65
χ2-test	12.28	7.34	8.19	10.13	9.13	9.20
p-value	[0.42]	[0.29]	[0.51]	[0.60]	[0.17]	[0.42]
r_f	0.007	0.01	0.011	-0.003	-0.019	-0.015
Wald test	**7.87**	**5.83**	**14.66**	**5.88**	0.23	0.68
p-value	[0.01]	[0.02]	[0.00]	[0.02]	[0.63]	[0.41]
Specification III						
β	**0.99**	**0.99**	**0.99**	**0.99**	**0.99**	**0.99**
	(0.001)	(0.002)	(0.001)	(0.005)	(0.005)	(0.004)
ρ	**1.56**	**2.28**	**1.76**	**-8.26**	**-7.36**	-5.38
	(0.19)	(0.91)	(0.61)	(2.62)	(3.32)	(4.07)
θ	**0.84**	**0.73**	**0.75**	**1.20**	0.93	**1.01**
	(0.02)	(0.05)	(0.03)	(0.11)	(0.53)	(0.40)
RRA	-0.31	-0.67	-0.32	10.90	7.86	6.42
χ2-test	14.31	11.68	13.07	14.98	13.27	14.16
p-value	[0.89]	[0.47]	[0.73]	[0.86]	[0.35]	[0.66]
r_f	0.011	0.008	0.009	0.002	0.0003	0.004
Wald test	**91.27**	**31.83**	**68.30**	3.42	0.02	0.0003
p-value	[0.00]	[0.00]	[0.00]	[0.06]	[0.90]	[0.99]

more volatile to cut through the minimum volatility bounds (MVB) and explain the variations in expected asset returns. The straight-bordered line is the Sharpe ratio of the aggregate

Table 3 Estimations with Abel's relative consumption model

Table reports the Abel (1990) catching up with the Joneses preference structure such that first panel reports the estimations for full sample and the 2nd panel for sub sample period. Specification I does the iterated GMM estimation using bond returns and return on aggregate stock index and specification II tests only with stock returns such that aggregate stock index ,SMB and HML as test assets. The specification III takes the bond returns and stock returns together in one cross section for model testing. In each panel against the subsequent instrument vector model parameters for time preference and risk preference are reported with notations β and α respectively. The standard errors for the model estimates are given in ().$\chi 2$-test is the Hansen's moment identification statistic reported with p-value in [].The predicted return on risk free rate with estimated parameters is reported against row $r_{f.}$ Wald test tests if H_o: α = 1 to draw on the fact if log normality for the model could be assumed reported with p-value in [].

	Panel I Full Sample			Panel II Sub Sample		
Specification I	INTI	INTII	INTIII	INTI	INTII	INTIII
β	**0.98**	**0.98**	**0.98**	**0.99**	**0.99**	**0.99**
	(0.002)	(0.007)	(0.004)	(0.004)	(0.003)	(0.003)
α	-0.59	-1.15	-0.36	**-3.90**	-1.93	-2.10
	(0.84)	(4.14)	(2.42)	(1.12)	(2.53)	(1.84)
$\chi 2$-test	13.98	7.06	8.56	10.39	9.17	9.84
p-value	[0.30]	[0.32]	[0.48]	[0.58]	[0.16]	[0.36]
r_f	0.018	0.02	0.018	0.014	0.013	0.013
Wald test	3.54	0.27	0.32	**19.08**	1.33	2.83
p-value	[0.06]	[0.60]	[0.57]	[0.00]	[0.25]	[0.09]
Specification II						
β	**0.99**	**0.98**	**0.98**	**0.99**	**0.99**	**0.99**
	(0.005)	(0.011)	(0.008)	(0.006)	(0.007)	(0.007)
α	-0.78	-1.58	-0.37	-3.20	-3.22	-3.22
	(1.80)	(6.25)	(5.61)	(1.45)	(4.10)	(4.10)
$\chi 2$-test	4.01	6.93	8.45	5.73	6.42	2.57
p-value	[0.98]	[0.86]	[0.49]	[0.93]	[0.38]	[0.98]
r_f	0.01	0.016	0.015	0.011	0.012	0.012
Wald test	0.98	0.17	0.06	3.01	1.06	1.06
p-value	[0.32]	[0.68]	[0.81]	[0.08]	[0.30]	[0.30]
Specification III						
β	**0.98**	**0.98**	**0.98**	**0.99**	**0.99**	**0.99**
	(0.002)	(0.008)	(0.005)	(0.003)	(0.005)	(0.003)
α	-1.47	-2.06	-1.92	**-2.42**	-4.32	-3.26
	(0.85)	(3.76)	(2.66)	(0.48)	(3.02)	(1.67)
$\chi 2$-test	14.60	12.24	13.35	11.10	9.98	10.49
p-value	[0.69]	[0.51]	[0.42]	[0.89]	[0.70]	[0.65]
r_f	0.016	0.016	0.016	0.012	0.012	0.012
Wald test	**8.47**	0.66	1.20	**49.96**	3.11	**6.53**
p-value	[0.00]	[0.42]	[0.27]	[0.00]	[0.08]	[0.01]

market index representing the equity premium (EP) zone.

All the model SDFs loci, for the subjective preference rate $\beta = 0.97$, are consistent with the EP zone, as shown in Figure 1 (a). The model SDFs also cut through the MVB, given perfect correlation, except for the EZ preferences. Nevertheless, the EZ preferences possess the largest SDF volatility among all. Still none have the variation to cut the squeezed bound to be consistent under truer SDF and asset return correlation estimate. Nonetheless, the alternative specifications reduce the risk aversion level for the Finnish agent more so than the standard CCAPM (Virk, 2012b) except for the Abel (1990) habit model. This fact is also evident from the lower curvature of the power utility-based SDF compared to the others.

The RRA for the standard power utility model should be around 7.9 and 9.1 to be consistent with the observed equity premium and asset return variations, respectively, for the Finnish market. The values for the risk parameter under Abel (1990) preferences are 10.3 and 12 respectively. The α values of 1.7 and 2.9 for the Constantinides (1990) model correspond to RRA values of 2.8 and 4.8 respectively to cut the EP zone and the MVB. Similarly, the curvature parameter values of 0.25 and 0.49 for CC SDF imply RRA estimates of 4.2 and 8.2 respectively. Cochrane (1997) argued that fewer solid reasons exist for objecting to the higher aversion to intertemporal substitution of agents than the higher risk aversion. Therefore, higher time preference under deterministic states of the world seems applicable for the historically higher risk free rate in the Finnish market relative to the U.S. and other developed stock markets (see Campbell, 1999).

The calibrations for the higher temporal impatience of the Finnish representative agent with $\beta = 0.95$ are presented in Figure 1 (b). The stretched locus for the model SDFs show that the noted ordinal structure for the RRA remains intact, as noted with the comparably lower temporal impatience. The calibrations with the allowance of stronger temporal preference show Constantinides (1990) and the CC internal habit models can even cut through the squeezed volatility bound. The other notable observation is the change in the location of the EZ-based SDF locus, which is now consistent with the EP region, MVB, and squeezed volatility bound. The EZ preferences have RRA estimates of 0.95, 4.85, and 5.85 to enter the corresponding regions respectively, which is fairly lower than the noted CCAPM RRA estimates.

The non-parametric calibrations display the importance of accounting investors' subjective time preferences. Assuming stronger time preferences improves the explanation of asset returns and reduces the severity of related 'puzzles' than imposing a tighter temporal behavior. The EZ and CC preferences are important examples in producing well-rounded results, under the plausible non-parametric values, for the Finnish representative agent. Moreover, the non- parametric estimates are consistent with the reported evidence in Cochrane and Hansen (1992) and Hyde et al. (2005).

Table 4 Estimations with Constantinides (1990) habit model

Table reports the Constantinides (1990) internal habit persistence model such that first panel reports the estimations for full sample and the 2nd panel for sub sample period. Specification I does the iterated GMM estimation using bond returns and return on aggregate stock index and specification II tests only with stock returns such that aggregate stock index ,SMB and HML as test assets. The specification III takes the bond returns and stock returns together in one cross section for model testing. The parameter θ determines the persistence of habit under the model. In each panel against the subsequent instrument vector model parameters for time preference and risk preference are reported with notations β and α respectively. The Table also computes the model risk aversion as described in section 2.3 reported against RRA. The standard errors for the model estimates are given in ().χ2-test is the Hansen's moment identification statistic reported with p-value in [].The predicted return on risk free rate with estimated parameters is reported against row r_f .Wald test tests if H_o: $\alpha = 1$ to draw on the fact if log normality for could be assumed for the model estimations reported with p-value in [].

	Panel I Full Sample			Panel II Sub Sample		
($\theta = 0.4$)	INTI	INTII	INTIII	INTI	INTII	INTIII
Specification I						
β	**1.44**	**1.19**	1.01	**0.48**	0.82	**0.41**
	(0.12)	(0.25)	(0.64)	(0.16)	(0.43)	(0.18)
α	**2.13**	**1.77**	**1.63**	**2.01**	**1.51**	**2.20**
	(0.22)	(0.26)	(0.33)	(0.40)	(0.53)	(0.57)
RRA	3.55	2.96	2.72	3.34	2.51	3.65
χ2-test	13.66	8.24	9.14	10.58	10.13	10.76
p-value	[0.32]	[0.22]	[0.42]	[0.57]	[0.12]	[0.29]
r_f	-0.37	-0.17	-0.01	0.74	0.20	0.89
Wald test	**27.49**	**8.93**	3.65	**6.55**	0.93	**4.36**
p-value	[0.00]	[0.00]	[0.06]	[0.01]	[0.34]	[0.04]
Specification II						
β	**1.54**	0.30	0.30	**0.14**	0.19	0.37
	(0.12)	(0.45)	(0.30)	(0.07)	(0.15)	(0.27)
α	**2.32**	2.66	2.62	**3.98**	**3.38**	**2.35**
	(0.24)	(2.36)	(1.51)	(0.94)	(1.34)	(1.06)
RRA	3.86	4.43	4.36	6.61	5.61	3.90
χ2-test	4.06	7.57	8.73	5.91	5.63	7.24
p-value	[0.98]	[0.27]	[0.46]	[0.92]	[0.47]	[0.61]
r_f	-0.43	1.21	1.19	1.97	1.64	1.00
Wald test	**28.95**	0.49	1.15	**10.04**	3.13	1.62
p-value	[0.00]	[0.48]	[0.28]	[0.00]	[0.08]	[0.20]
Specification III						
β	**1.53**	0.95	**0.96**	**0.57**	**0.41**	**0.60**
	(0.04)	(0.48)	(0.44)	(0.20)	(0.13)	(0.22)
α	**2.31**	**1.61**	**1.62**	**1.83**	**2.21**	**1.79**
	(0.09)	(0.15)	(0.15)	(0.34)	(0.43)	(0.32)
RRA	3.85	2.69	2.69	3.04	3.67	2.97
χ2-test	15.01	11.09	12.64	9.51	10.83	10.17
p-value	[0.66]	[0.60]	[0.48]	[0.95]	[0.63]	[0.68]
r_f	-0.43	0.05	0.04	0.57	0.89	0.52
Wald test	**23.16**	**13.14**	**27.12**	**6.06**	**7.91**	**6.14**
p-value	[0.00]	[0.00]	[0.00]	[0.01]	[0.00]	[0.01]

5.2. Model estimations

The estimations are presented with the model-wise hypothesis testing on the plausibility of model parameters and local restrictions on the models, such as differentiating them from

Table 4 (Continued).

	Panel I Full Sample			Panel Sub Sample		
	INTI	INTII	INTIII	INTI	INTII	INTIII
Specification I						
β	**0.35**	0.02	0.83	**0.28**	**1.16**	0.24
	(0.17)	(0.03)	(0.49)	(0.13)	(0.24)	(0.16)
α	**1.45**	**3.99**	**1.32**	**1.62**	**1.14**	**1.74**
	(0.30)	(1.48)	(0.39)	(0.35)	(0.66)	(0.54)
RRA	3.62	9.95	3.29	4.03	2.83	4.33
$\chi2$-test	12.92	9.68	9.08	10.63	11.98	10.73
p-value	[0.38]	[0.14]	[0.43]	[0.56]	[0.06]	[0.29]
r_f	1.03	3.73	0.18	1.27	-0.15	1.42
Wald test	2.26	**4.06**	0.68	3.23	0.04	1.91
p-value	[0.13]	[0.04]	[0.41]	[0.07]	[0.83]	[0.17]
Specification II						
β	-	0.20	0.19	-	0.17	0.30
	-	(0.48)	(0.27)	-	(0.16)	(0.32)
α	-	1.91	1.96	-	**2.07**	**1.57**
	-	(2.15)	(1.30)	-	(0.91)	(0.78)
RRA	-	4.77	4.89	-	5.14	3.90
$\chi2$-test	-	12.92	8.66	-	5.69	7.11
p-value	-	[0.27]	[0.47]	-	[0.46]	[0.63]
r_f	-	1.62	1.67	-	1.78	1.20
Wald test	-	0.18	0.55	-	1.37	0.54
p-value	-	[0.67]	[0.46]	-	[0.24]	[0.46]
Specification III						
β	-	0.67	**0.78**	**0.40**	**0.29**	**0.29**
	-	(0.63)	(0.37)	(0.21)	(0.13)	(0.13)
α	-	**1.24**	**1.28**	**1.39**	**1.62**	**1.62**
	-	(0.10)	(0.22)	(0.30)	(0.35)	(0.35)
RRA	-	3.10	3.20	3.45	4.02	4.02
$\chi2$-test	-		12.49	9.23	10.72	10.72
p-value	-	[0.60]	[0.49]	[0.95]	[0.63]	[0.63]
r_f	-	0.40	0.25	0.92	1.26	1.26
Wald test	-	**6.34**	1.74	1.73	3.11	3.11
p-value	-	[0.01]	[0.10]	[0.19]	[0.08]	[0.08]

CRRA preferences or testing for the log normality. The model-based predictions for the risk free rate are also expounded in reporting the success of the model SDFs. We close this section by comparing the competing model HJ-distance. The model implications on stock and bond return variability are analyzed across three test asset specifications. Specification I includes returns on the aggregate stock market, proxy risk free rate, and a government bond with a 10-year maturity as test assets. Specification II includes returns on stock portfolios, such as the aggregate market index and Fama and French (1993) SMB and HML portfolio strategies. Specification III tests the real strength of the estimated utility modifications to explain all the assets together.

Table 2 reports the model estimations for the EZ (1991) state non-separable preferences. The model moment restrictions cannot be rejected by the Hansen's (1982) test. The estimate for ρ parameter, which reflects substitution, is quite unstable across the full period specifications. The large estimate (with INTII and INTIII) for bond returns signifies higher aversion to switch across periods. However, the substitution effect is comparably lower for specification II (with INTII and INTIII), whereas the model implied EIS is low when stock and bond returns are taken together. The subsample estimations reveal preferences for the lower implied EIS. The implied risk aversion parameter in the full sample is low, and estimations providing unreasonable risk preferences in specification III are nonconforming to the evidence in EZ (1991). The implausible risk implications are limited to the full sample estimations only, though the implied RA estimates are economically sensible in the recovery period across all the specifications.

The notable consistency with the EZ estimations is for the time preference parameter δ, which is always greater than zero. The plausible identification of the temporal preference keeps the discount factor below 1 across all the samples and specifications, an improvement on the original study for the U.S. market. The noted economic structure projects positive return on the risk free asset, whereas the impatience estimates in the EZ (1991) provide counterfactual predictions under deterministic states of the world. The subjective time behavior for the Finnish agent is consistent with the evidence in Viitanen (2004) for the samples accounting for financial liberalization (post 1987).

Generally, the ratio parameter θ maintains earlier resolution of uncertainty in the full sample consistent with Viitanen (2004), whereas the stable period estimations signify late resolution of uncertainty. The variations in the resolution of uncertainty across samples signify a conditional relationship given representative agents' EIS in consumption between today and tomorrow. The Finnish agents in the recovery period are less averse to substituting consumption across periods than the full sample, given estimated risk preferences. The effect of the relative disliking of intertemporal fluctuations in consumption is also manifested in the higher predicted returns on the risk free rate in the full sample estimations rather than in the recovery period estimations.

Besides, the agents are more risk averse in the recovery period relative to the full period. The (plausible) RRA estimates in the full sample are below 2 compared to the sufficiently

Table 5 Campbell and Cochrane Surplus Consumption Ratio Model

Table reports the Campbell and Cochrane (1999) external habit persistence model such that first panel reports the estimations for full sample and the second panel for sub sample period. Specification I does the iterated GMM estimation using bond returns and return on aggregate stock index and specification II tests only with stock returns such that aggregate stock index, SMB and HML as test assets. The specification III takes the bond returns and stock returns together in one cross section for model testing. In each panel against the subsequent instrument vector model parameters for time preference and risk preference are reported with notations β and α respectively. The Table also computes the model risk aversion as described in section 2.4 reported against row heading RRA. The standard errors for the model estimates are given in ().$\chi2$-test is the Hansen's moment identification statistic reported with p-value in [].The predicted return on risk free rate with estimated parameters is reported against row r_f.

	Panel I Full Sample			Panel II Sub Sample		
Specification I	INTI	INTII	INTIII	INTI	INTII	INTIII
β	**0.98**	**0.97**	**0.98**	**0.98**	**0.99**	**0.97**
	(0.001)	(0.026)	(0.002)	(0.002)	(0.001)	(0.008)
α	**0.08**	0.64	**0.11**	**0.08**	0.10	0.20
	(0.02)	(1.13)	(0.05)	(0.02)	(0.15)	(0.15)
RRA	1.36	10.36	1.72	3.54	4.09	8.58
$\chi2$-test	12.96	7.78	8.54	10.30	9.02	9.65
p-value	[0.45]	[0.35]	[0.58]	[0.67]	[0.25]	[0.46]
r_f	0.016	0.022	0.016	0.020	0.014	0.024
Specification II						
β	**0.99**	**0.91**	**0.99**	**0.99**	**0.98**	**0.98**
	(0.004)	(0.059)	(0.003)	(0.005)	(0.010)	(0.004)
α	0.03	**1.65**	0.18	**0.23**	0.71	0.14
	(0.06)	(0.79)	(0.10)	(0.07)	(0.78)	(0.14)
RRA	0.48	26.63	2.85	9.66	30.06	5.87
$\chi2$-test	4.00	6.27	9.43	6.27	5.78	7.45
p-value	[0.99]	[0.51]	[0.49]	[0.94]	[0.57]	[0.84]
r_f	0.008	0.022	0.013	0.007	-0.026	0.015
Specification III						
β	**0.99**	**0.95**	**0.99**	**0.99**	**0.98**	**0.99**
	(0.001)	(0.020)	(0.001)	(0.001)	(0.003)	(0.001)
α	**0.04**	**1.12**	**0.09**	**0.16**	0.46	**0.15**
	(0.01)	(0.52)	(0.02)	(0.03)	(0.45)	(0.06)
RRA	0.59	16.58	1.33	4.73	13.72	4.60
$\chi2$-test	14.79	10.77	10.77	11.10	9.69	10.37
p-value	[0.68]	[0.63]	[0.63]	[0.89]	[0.72]	[0.66]
r_f	0.013	0.017	0.013	0.011	-0.005	0.013

higher values for RRA in the subsample estimations. The full sample estimations reject the expected utility preferences. The model predicted return on risk free rate is precise in the full sample compared to the recovery periods' smooth predictions. On average, the full sample predictions misprice the quarterly risk free rate by 40bps in specification I, whereas mispricing is only 10bps for specifications II and III. In comparison, the subsample projections misprice the quarterly return in specifications I, II, and III by an average 60bps, 210bps, and 60bps respectively. The EZ (1991) preferences for the recovery period predict returns on the risk free rate negatively in specification II. The EZ (1991) estimations provide varying implications for investors' temporal subjectivity, risk preferences, and elasticity to substitute consumption across periods. Nevertheless, the suitability of recursive preferences in the full sample is backed by better risk free predictions, whereas the lesser aversion to intertemporal substitution leads to implausibly smooth predictions for the return on risk free rate. The plausible estimates for the substitution effect ρ and risk preferences are sensitive to the selection of test assets in the full sample.

The estimations with the "catching up with Joneses" preferential structure are reported in Table 3.Abel's external habit model faces great difficulty in producing plausible risk preferences across all the reported estimations. The model estimations display marginally higher subjective time preference yet remain plausible in the full sample estimates. The stronger time preference leads to a larger prediction for the return on risk free rates in the full sample than the ex-recession period. The model forecasts for the risk free rate are above the observed rate by 100bps, 40bps, and 30bps for specification I, II, and III respectively. The corresponding mispricing for the sub sample estimations is 40bps, 30bps, and 30bps. The log normality assumption is rejected frequently for the specified consumption growth and external habit levels.

The results with the Constantinides (1990) habit model are presented in Table 4 for varying degrees of subsistence level. The implications for investor impatience parameters are grossly implausible. The impatience estimate implies agents are either more interested in shifting their consumption to the next period, the case when the subjective preferences are estimated greater than 1, or they are highly impatient to consume now such that the β values are sufficiently below 1. The first case is economically inconsistent and corresponds to negative interest rate predictions. The second case requires high interest rates for inducing substitution, which is nonviable. Moreover, the estimates for the subjective preferences are not consistent with the tighter estimates we obtained from other specified model estimations.

The subsample estimations are also marred with extreme time preferences for the Finnish representative agent, such that the Finnish agent's eagerness to consume today corresponds to high interest rate predictions, despite the degree of habit persistence, when compared with the full period estimations. The plausible temporal preferences are recorded with $\theta = 0.4$ in specification III with INTII and INTIII, although these preferences are still higher than the estimates from other specified models. The local curvature parameter α corresponds to plausible and significant risk aversion values. However, the extreme impatience estimates devastate the possibility for a coherent reconciliation with the observed facts. The Wald test

Table 6 Predictability regressions

	Panel A			Panel B			Panel C	
	$r^e_{m,t+1}$	$r^e_{LGB,t+1}$		$r^e_{m,t+1}$	$r^e_{LGB,t+1}$		$r^e_{m,t+1}$	$r^e_{LGB,t+1}$
β_s	-0.02	**0.01**	β_s	-0.01	**0.01**	β_s	0.004	**0.01**
t_{NW}	(-0.23)	(2.71)	t_{NW}	(-0.07)	(2.63)	t_{NW}	(0.04)	3.04
β_{ts}	1.96	**0.14**	β_{pd}	0.21	0.01	β_{per}	-0.75	0.02
t_{NW}	(1.65)	(6.83)	t_{NW}	(0.74)	(1.60)	t_{NW}	(-0.81)	0.64
$\bar{R}^2$	0.03	0.43		0.01	0.20		0.01	0.15

is for the case if $\alpha = 1$ under the null hypothesis and tests for the log normality of the data, an assumption under which we approximated the predictions for the risk free rate. The log normality null for the consumption data with $\theta = 0.4$ is rejected more frequently than when we specify habit persistence for $\theta = 0.6$.

The model estimations for the surplus consumption ratio model are presented in Table 5.[9] The subsample estimations report higher risk aversion values than the full sample, consistent with the EZ (1991) estimations across samples. The estimations provide precise power coefficient estimates across both the samples using INTI. Generally, the steady state risk aversion is higher with INTII than the other instrument vectors. The surplus consumption specification also predicts the return on risk free security with larger forecasting errors in the full period than the observed rate. Nevertheless, the interest rate projections in the full sample reported with INTI and INTIII are better than INTII. The reported mispricing across the specifications on average is 110bps, 50bps, and 40bps respectively.

The predicted returns with CC (1999) are reasonably lower than the estimated habit persistence models of Abel (1990) and Constantinides (1990). The convergence of power coefficient to positive values is documented to be consistent with Engsted et al. (2010). The CC (1999) estimated model parameters are economically more plausible than the other tested utility specifications inclusive of the EZ (1991) recursive preferences. The specification imparts theoretical structure on the Finnish representative agents' subjective time preferences and risk aversion. Moreover, the forecasting ability of the CC (1999) specification for the return on risk free rate improves in the subsample estimations.

[9] The statistics in the calibration of the surplus consumption ratio are available upon request. The surplus consumption ratio across the samples is time varying and the habit is closer to the consumption levels in the subsample than those in the full sample calibrations.

The surplus consumption ratio has noteworthy implications for the asset return predictability.[10] Therefore, like Engsted et al. (2010), we also test for the return predictability prowess of the log process for the Finnish stock and bond returns. We run bi-variate regressions including other known predictability variables such as pd, pe ratios and term structure of interest rates ts along with surplus consumption ratio. All the variables in the predictability regressions are log linearized. We estimate the following regressions $r^e_{m,t+1} = \alpha + \beta_s s_t + \beta_x x_t + e_{t+1}$ and $r^e_{LGB,t+1} = \alpha + \beta_s s_t + \beta_x x_t + e_{t+1}$ for the excess stock and bond return predictability. The explanatory variable x_t, corresponds to mentioned predictors accordingly. The results show none of the lagged predictors, including log surplus consumption ratios, have any considerable predictability for the excess premium on the stocks, as reported in Table 6. However, the surplus consumption ratio has considerable forecasting ability for the excess bond premium. The lagged surplus ratio remains significant even in the presence of term structure on interest rates. The better predictability for the excess bond premia compared to the excess equity premia is also evident in the comparable $\bar{R}^2$ values.

Hansen and Jagannathan (1997) pricing errors for all the estimations are reported in Table 7. The EZ (1991), Abel (1990), and CC (1999) markedly improve on the pricing errors from standard power utility based CCAPM for the Finnish market reported in Virk (2012b).[11] The Constantinides (1990) model faces larger difficulty in pricing the assets compared to the other model pricing kernels. The mispricing is aggravated when the habit level is increased to 0.6, an observation generalized across all specifications and samples. The dismal performance of the Constantinides (1990) is a further testament to the empirical failure to explain Finnish temporal and risk preferences. The HJ mispricing is larger when the models are estimated with the bond returns – that is, EZ estimations in specification I only or when they are taken together with the stock returns as is the case in specifications I and III for the

[10] The CC (1999) model captures the variations in expected stock returns with closeness of consumption to habit persistence, that is, surplus consumption ratio. Therefore, the model implies that for periods with lower surplus consumption ratio, agents are more averse (counter cyclic) and require higher (expected) returns for holding risky assets and vice versa. Therefore, the surplus consumption should capture the information to predict future paths of time-varying expected returns.

[11] The comparative HJ-distance in Virk (2012b) with the standard power utility model corresponding to estimations specifications I and III in this study are on average 42 percent and 118 percent, whereas, the mispricing for the subsample is 58 percent and 138 percent. To compare the specification II CCAPM implied HJ specification errors and the specifications tested in this paper, we estimated the CCAPM HJ pricing errors. The pricing errors in the full period and the subsample are 0.12, 0.11, 0.11, and 0.09, 0.08, and 0.1 respectively with the instruments used in this study. The CCAPM implied pricing errors tend to outperform the habit models in the stable period whereas the Campbell and Cochrane (1999) model does marginally better than the power utility model in the full period.

Table 7 HJ-distances for modified specifications

HJ-distance is the Hansen and Jagannathan (1997) distance measure computed as in equation (15) and σ(HJ) reports the asymptotic standard errors given in () for the computed HJ-distance as in (16) . The HJ-distances for each preference specifications are estimated using the inverse of the covariance matrix of the test assets in the respective specification.

		Panel I Full Sample			Panel II Sub Sample		
Specification I		INTI	INTII	INTIII	INTI	INTII	INTIII
Epstein and Zin Model							
	$HJ-distance$	0.45	0.45	0.45	1.17	1.19	1.18
	$\sigma(HJ)$	(0.24)	(0.25)	(0.24)	(0.30)	(0.30)	(0.30)
Abel Model							
	$HJ-distance$	0.45	0.45	0.46	1.21	1.21	1.21
	$\sigma(HJ)$	(0.24)	(0.24)	(0.24)	(0.29)	(0.29)	(0.29)
Constantinides Model							
$\theta = 0.4$	$HJ-distance$	0.63	0.49	0.45	1.35	1.25	1.37
	$\sigma(HJ)$	(0.17)	(0.23)	(0.25)	(0.27)	(0.29)	(0.26)
$\theta = 0.6$	$HJ-distance$	0.78	1.07	0.48	1.44	1.24	1.46
	$\sigma(HJ)$	(0.15)	(0.11)	(0.24)	(0.25)	(0.29)	(0.25)
Campbell and Cochrane Model							
	$HJ-distance$	0.45	0.40	0.45	1.20	1.20	1.20
	$\sigma(HJ)$	(0.24)	(0.26)	(0.24)	(0.29)	(0.29)	(0.29)
Specification II							
Epstein and Zin Model							
	$HJ-distance$	0.02	0.01	0.01	0.03	0.01	0.01
	$\sigma(HJ)$	(0.09)	(0.09)	(0.08)	(0.12)	(0.11)	(0.11)
Abel Model							
	$HJ-distance$	0.11	0.12	0.12	0.09	0.15	0.10
	$\sigma(HJ)$	(0.12)	(0.14)	(0.13)	(0.14)	(0.15)	(0.14)
Constantinides Model							
$\theta = 0.4$	$HJ-distance$	0.57	0.70	0.69	0.85	0.80	0.63
	$\sigma(HJ)$	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
$\theta = 0.6$	$HJ-distance$	-	0.80	0.81	-	0.82	0.69
	$\sigma(HJ)$	-	(0.02)	(0.02)	-	(0.02)	(0.02)
Campbell and Cochrane Model							
	$HJ-distance$	0.12	0.09	0.11	0.12	0.11	0.12
	$\sigma(HJ)$	(0.12)	(0.13)	(0.12)	(0.13)	(0.14)	(0.13)
Specification III							
Epstein and Zin Model							
	$HJ-distance$	0.46	0.46	0.46	1.15	1.14	1.15
	$\sigma(HJ)$	(0.24)	(0.23)	(0.23)	(0.32)	(0.32)	(0.32)
Abel Model							
	$HJ-distance$	0.46	0.46	0.46	1.21	1.21	1.21
	$\sigma(HJ)$	(0.24)	(0.24)	(0.24)	(0.29)	(0.29)	(0.29)
Constantinides Model							
$\theta = 0.4$	$HJ-distance$	0.71	0.46	0.46	1.32	1.38	1.32
	$\sigma(HJ)$	(0.15)	(0.24)	(0.24)	(0.27)	(0.26)	(0.27)
$\theta = 0.6$	$HJ-distance$	-	0.56	0.50	1.39	1.44	1.44
	$\sigma(HJ)$	-	(0.20)	(0.22)	(0.26)	(0.25)	(0.25)
Campbell and Cochrane Model							
	$HJ-distance$	0.47	0.36	0.46	1.20	1.19	1.20
	$\sigma(HJ)$	(0.24)	(0.25)	(0.24)	(0.29)	(0.30)	(0.29)

Table 8 Beta risk vs. Consumption risk

The Table reports the results from the Chen et al. (1986) based test which specifies CAPM based beta risk to be zero and consumption based model risks to be positive under the null hypothesis. β_m ,β_{CCAPM} , β_{EZ} and β_{CC} correspond to the log linearized CAPM beta, standard CCAPM, EZ (1991) and Campbell and Cochrane (1999) based consumption risks. The suffixation of consumption risks with cg and dg highlights the estimation of consumption risks with stocks as claim to aggregate consumption or as a claim to (volatile) dividends as in the Lucas (1978) model. The significant regression estimates at 5% and below significance are provided in bold fonts. The Newey and West (1987) corrected t-values are reported in () Additionally, $\overline{R}^2$ values for the Chen et al.(1986) regressions for all the assets are reported.

	Panel I: regressions with cg			Panel II: regressions with dg		
	β_m	β_{CCAPM_cg}	$\overline{R}^2$	β_m	β_{CCAPM_dg}	$\overline{R}^2$
r^e_{sl}	**0.48**	0.88	0.38	**0.49**	**-0.21**	0.42
	(3.81)	(1.10)		(4.01)	(-2.46)	
r^e_{sm}	**0.47**	0.82	0.28	**0.48**	**-0.24**	0.37
	(5.50)	(1.04)		(5.77)	(-2.06)	
r^e_{sh}	**0.54**	-1.38	0.41	**0.53**	-0.01	0.39
	(6.04)	(-1.07)		(6.62)	(-0.08)	
r^e_{bl}	**0.93**	0.17	0.80	0.93	-0.09	0.81
	(12.65)	(0.29)		(13.01)	(-1.57)	
r^e_{bm}	**0.64**	-0.51	0.55	0.63	-0.12	0.55
	(4.07)	(-0.62)		(4.26)	(-1.62)	
r^e_{bh}	**0.51**	0.12	0.43	0.56	-0.16	0.46
	(4.33)	(0.51)		(4.66)	(-1.68)	

	β_m	β_{EZ_cg}	$\overline{R}^2$	β_m	β_{EZ_dg}	$\overline{R}^2$
r^e_{sl}	0.19	1.17	0.39	**0.53**	-0.01	0.39
	(0.57)	(1.10)		(7.06)	(-0.08)	
r^e_{sm}	0.20	1.09	0.33	**0.56**	**-0.32**	0.37
	(0.72)	(1.04)		(6.58)	(-2.06)	
r^e_{sh}	**1.00**	-1.83	0.41	**0.56**	**-0.28**	0.42
	(2.05)	(-1.07)		(4.95)	(-2.46)	
r^e_{bl}	**0.88**	0.22	0.80	**0.96**	-0.11	0.80
	(3.99)	(0.29)		(14.89)	(-1.57)	
r^e_{bm}	**0.81**	-0.68	0.55	**0.67**	0.15	0.56
	(2.14)	(-0.62)		(4.56)	(-1.62)	
r^e_{bh}	0.26	1.01	0.43	**0.57**	-0.21	0.45
	(0.47)	(0.51)		(5.19)	(-1.68)	

	β_m	β_{CC_cg}	$\overline{R}^2$	β_m	β_{CC_dg}	$\overline{R}^2$
r^e_{sl}	**0.50**	1.53	0.38	**0.49**	0.12	0.38
	(3.87)	(0.98)		(3.87)	(0.32)	
r^e_{sm}	**0.48**	-0.15	0.32	**0.48**	**1.03**	0.37
	(5.44)	(-0.09)		(5.30)	(2.91)	
r^e_{sh}	**0.54**	2.29	0.40	**0.53**	-0.02	0.40
	(6.54)	(0.63)		(6.61)	(-0.06)	
r^e_{bl}	**0.95**	**4.64**	0.82	**0.93**	**0.55**	0.81
	(13.56)	(4.28)		(13.77)	(2.79)	
r^e_{bm}	**0.63**	-1.06	0.55	**0.63**	0.15	0.55
	(4.17)	(-0.54)		(4.17)	(0.48)	
r^e_{bh}	**0.55**	**8.97**	0.53	**0.21**	0.12	0.42
	(4.75)	(5.56)		(0.92)	(0.23)	

habit models.

The noted simplification shows that alternative preferences have more difficulty than the stock returns in corroborating the bond returns. Nonetheless, the surplus consumption ratio model suppresses HJ-distance better for the specifications including bond returns and, arguably, for the reported ability to predict bond returns. The external habit model of Abel (1990) has a similar level of mispricing as the CC (1999) across specifications. However, the inability to give plausible risk preferences, as reported in Table 3, undermines the overall model performance to fit with asset pricing data.

EZ (1991) prices the stock returns in specification II with sufficiently lower errors compared to the CC (1999) habit model across both samples. The success of EZ (1991) preferences for the stock returns demonstrates Cochrane and Hansens' (1992) point that the model IMRS has larger variability for including return on the market index.[12] In contrast, CC (1999) produces the smallest HJ-distance with INTII for the full sample in specifications I and III amongst the specified models. Otherwise the pricing errors from CC (1999) and EZ (1991) are nearly identical. Moreover, the model estimations for the habit persistence specifications also explain the return on the market index whereas, the EZ (1991) preferences take it exogenously given. Therefore, the results support the CC (1999) surplus consumption ratio model for its greater ability to suppress mispricing relative to other preferential modifications.

The HJ-distances in the subsample estimations, for specifications I and III, are substantially large to be anything but consistent with the asset returns; Virk (2012b) also documented the same for the standard CCAPM.

5.3. Diagnostic analysis

We take the successful modifications, along with the standard CCAPM model, for further investigation to check the consistency of the main results. Thus, the CC (1999) surplus consumption ratio model and EZ (1991) preferences' (despite structural constraints) are selected.

[12] The construction constraint of the EZ (1991) model SDF factually limits the model comparison between the recursive utility specification and the habit persistence models for using different HJ-weighting matrices. This structural limitation of the EZ (1991) preferences perturbs the constancy of weighting matrix.

In the first test, we follow the motivation of Mankiw and Shapiro (1986) that consumption risk should theoretically be a better measure of systematic risk than the CAPM beta.[13] Therefore, the proposed test in Chen et al. (1986) is employed for reporting whether the CAPM risk (Sharpe, 1964; Lintner, 1965) or CCAPM based consumption risk is significant to influence the cross-section of assets. All model SDFs are log linear from herein for reporting comparable joint hypotheses testing with the CAPM model.[14] We construct six benchmark size and book-to-market (size-BM) sorted Fama and French (1993) portfolios for the Finnish market from 1990:01 onwards.

The Chen et al. (1986) based test suggests a regression of type $r^e_{j,t+1} = \beta_{m,j} r^e_{m,t+1} + \beta_{x,j} x_{t+1} + e_{t+1}$, such that $\beta_{m,j}$ and $\beta_{x,j}$ correspond to market model beta risk and consumption-based model beta risk respectively for asset j. The latter represents standard CCAPM, EZ (1991), or CC (1999) model-based beta risks for asset j accordingly. The hypotheses testing for the consumption-based models specify $\beta_{m,j} = 0$ and $\beta_{x,j} > 0$ under the joint null hypothesis. The significance of both the beta risks is tested with Newey and West (1987) based t-ratios. Moreover, following Lucas (1978), the estimations are repeated with dividend growth, assuming consumption is contingent on dividends.

The results in the first vertical panel show the null hypothesis for the CAPM-based beta risk is rejected across all the assets. The beta risk is always positive, whereas CCAPM-based consumption risk is implausibly negative for the portfolios sh and bm. Moreover, none of the consumption risk is significant to explain the time series variations in the test portfolios. The performance of CAPM beta is less convincing when taken against EZ (1991) consumption beta such that the former is only significantly positive for sm, sh, and bl portfolios. However, EZ (1991) linearized SDF is positive for portfolios sl, sm, bl, and bh although insignificantly estimated. The CC (1999) based consumption risk is significantly positive for portfolios bl and bh, a feat not achieved by other consumption-based model risks. Similar to

[13] In theory, the risk under the consumption-based models has more encompassing implications for assigning payoffs to the stocks. First, it incorporates the agents' decision-making preferences across periods. Second, the measure governing investor portfolio decision making in different states of the world implicitly incorporates additional sources of wealth compared to the wealth from only the equity market.

[14] For example, we take the log of the consumption-based model IMRS such that the standard power utility based SDF, that is, $\beta(\frac{C_{t+1}}{C_t})^{-\alpha}$ is transformed to $-\alpha\beta\left(\frac{c_{t+1}}{c_t}\right) = \beta_{CCAPM}\left(\frac{c_{t+1}}{c_t}\right)$. The exogenous variables for log linearization are reported in lower case. Moreover, β_{CCAPM} in this context signifies both the substitution effect and investor risk aversion. Similar transformations are also employed for EZ (1991) and CC (1999) model SDFs. The EZ (1991) model IMRS is aggregated across expected utility and non-expected utility components for $\theta = 0.75$ with $\rho = 3$, in consideration to the available results in Table 2 for Finnish market.

Table 9 Linear SDF factor sensitivities and cross-sectional prices of model risks

The SDF-cross-sectional regressions are reported for the factor sensitivities of the linearized consumption based models and correspondently the prices of consumption risks for CCAPM, EZ (1991) and Campbell and Cochrane (CC, 1999) model specifications. The last of part of the table reports the estimations for unconditional CAPM. The cross-sectional SDF estimations are done with the two-step GMM using Hansens' (1982) optimal weighting matrix using Newey and West (1987) estimator with Bartlett kernel. The estimations use the returns on six size-BM portfolios along with return on the risk-free rate for the tractability of SDF mean across all the models. The a_x, and b_x are SDF sensitivities for intercept and model factors, whereas and λ_x is the price of risk for the linearized SDF under respective model. The suffixation of cg and dg in panel I and panel II highlights the estimation of consumption risks with stocks as claim to aggregate consumption the standard practice or as a claim to dividend streams as in the Lucas (1978) model. The regression estimates at 5% and below significance are provided in bold fonts. The Newey and West (1987) corrected t-values are reported in (). Hansens' (1982) over identification test is given in front of row heading TJ_T with corresponding p-values in [].

	Panel I with cg		Panel II with dg	
CCAPM				
a_{CCAPM}	**1.01**	(80.87)	**1.04**	(9.67)
b_{CCAPM}	**-10.43**	(-1.74)	**10.05**	(9.69)
λ_{CCAPM}	0.002	(1.69)	0.19	(0.43)
TJ_T	13.00	[0.45]	9.93	[0.70]
EZ Model with $\theta = 0.75$				
a_{EZ}	**1.00**	(99.70)	**0.99**	(4.36)
b_{EZ}	**-4.62**	(-87.35)	**9.88**	(14.36)
λ_{EZ}	**0.01**	(56.75)	**-2.71**	(-3.37)
TJ_T	12.45	[0.49	11.42	[0.58]
CC Model				
a_{CC}	**0.98**	(188.89)	**0.96**	(77.14)
b_{CC}	**-14.62**	(-1.71)	**-6.04**	(-72.36)
λ_{CC}	0.001	(1.65)	**0.01**	(5.84)
TJ_T	11.94	[0.53]	12.82	[0.46]
		CAPM		
	a_m	**1.00**	(72.78)	
	b_m	**-1.36**	(-9.56)	
	λ_m	**0.06**	(8.29)	
	TJ_T	12.45	[0.49]	

other consumption-based model SDFs, the model pricing kernel also fails to significantly influence the time series variations of the test portfolios relative to CAPM risk.

The regressions replacing consumption growth with dividend growth (panel II of Table 8) fare no better for the CCAPM; in fact, they are even poorer such that the point estimates for the consumption risk are negative. Similar performance is reported with the EZ (1991) based risk modeled with dividend growth. We get the most improved performance by the CC (1999) based SDF among the consumption-based models while using dividend growth. The consumption risk is significantly priced for sm and bl, whereas for the remaining portfolios it is positive except for portfolio sh. Nonetheless, the CAPM beta risk maintains supremacy compared to all the tested consumption risks.

In the second robustness check, we test for the model-based linear SDF factor sensitivities and the cross-sectional price of the risk estimate. The estimation procedure approximates all the moments jointly, as suggested in Cochrane (2005), to adjust for generated regressors problem faced in otherwise standard, two pass, cross-sectional regressions. The significance of SDF factor sensitivities and the cross-sectional price of risk estimate is reported for 5 percent confidence GMM t-values. The GMM t-values are corrected for heteroskedasticity and autocorrelation in the model errors with the Newey and West (1987) method. The unconditional moments for each model at time $t + 1$ are such that

$$u = \begin{pmatrix} (a + bx)r_j^e \\ , \\ (r_j^e - \lambda_x \beta_j) \end{pmatrix} \otimes Z \qquad (18)$$

where $\boldsymbol{u}$ is the vector of GMM errors, which are minimized to zero from the sample counterpart of the imposed model moments. The instrument vector across all the tested models is $(1\ r_{m,t-1}\ r_{m,t-2})$. a and b are the factor sensitivities for the respective model SDF; for example, x is the log consumption growth for the standard CCAPM, whereas β_j is the model risk, and λ is the estimated unconditional price of risk for the risk factor x.

The results based upon the above moments, across the cross-section of six size-BM portfolios along with the return on risk free security, are reported in Table 9. The return on risk free security is included for the tractability of the unconditional SDF mean. The moment restrictions from all the SDF cross-sectional regressions cannot be rejected with Hansens' (1982) test at 5 percent conventional p-values. The test suggests standard CCAPM consumption risk factor sensitivity, and the cross-sectional prices of model risk are insignificantly plausible. However, the EZ (1991) factor risk significantly influences the linear pricing kernel and also has plausible price of risk. The CC (1999) model estimations also imply plausible estimates for the factor sensitivity and factor risk premium, however insignificant at 5 percent critical values. The estimations reported in panel B with dividend growth do not yield encouraging results for CCAPM and the EZ (1991) recursive utility model. The results for the EZ (1991) model SDF using dividend growth have a rather dismal performance for model factor risk sensitivity and price of risk estimates, both of which are theoretically implausible.

Similar to its better performance in Table 8 with dividend growth, the CC (1999) model produces plausible factor risk sensitivity and price of risk estimate. Additionally, both are significant at 1 percent critical GMM t-values. The reported result for the CAPM based-SDF provides model-consistent estimates for the model factor risk and price of risk at 1 percent significant GMM t-values.

5. Conclusions

The implications of the consumption-based equilibrium models have taken many routes since the seminal work of Mehra and Prescott (1985). These widespread developments has focused on modifying preferences central to standard models, heterogeneity of agents and uninsurable risk, survivorship bias, and a number of constraints related to market frictions such as borrowing costs, liquidity, transaction costs, and taxation. The general test laboratory for these theoretical and empirical developments has been U.S. stock market. A recent surge in reporting the evidence for the modified preferences to the standard power utility specification (CCAPM) from non-U.S. financial markets has been observed (e.g., Hyde et al., 2005; Engsted & Møller, 2010; Engsted et al., 2010). Generally, the evidence has focused on comparing the performance of the leading habit formation model of CC (1999) with the standard power utility based model.

This study provides evidence from the Finnish market while estimating a number of proposed preferential modifications, such as EZ (1991), Abel (1990), and Constantinides (1990) along with CC (1999), for a complete comparison. The Finnish stock market has observed large equity premia in the sample although with comparably low reward-to-risk ratios. This idiosyncrasy puts the Finnish market at the margin of capitalized and European economies and provides an added dimension to the earlier reported international evidence. The non-parametric testing with Hansen and Jagannathan's (1991) methodology shows that the EZ (1991), Constantinides (1990), and CC (1999) models generate mean and volatility pairs to explain variations in asset returns. The better performance of Constantinides (1990) is not transformed to generate plausible parametric solutions to describe aggregate equity premia and other related facts. The EZ (1991) preferences provide mixed evidence and show that Finnish representative agents have varying risk tolerances across the samples. Importantly, CC (1999) estimations also provide similar evidence with lower risk aversion in the full sample than in the recovery period estimations. CC (1999) is the only model providing economically sensible parameterization across all the samples and specifications.

Overall, the analytic solutions and parametric estimations for the better performing models show stronger time preference behavior (such as disliking temporal shifts in consumptions) reduces the severity of the puzzles.

In terms of HJ (1997), distance measure EZ (1991) gives the lowest pricing errors for specifications taking only stock returns. CC (1999) produces the lowest distance for specifications including the bond returns with the stock returns among all. The better performance of EZ (1991) for stocks should not come as a surprise for using the return on aggregate market index in EZ’s specification IMRS. Thus, we conclude the CC model performs the best among the tested consumption-based equilibrium models. The surplus consumption ratio significantly predicts the bond returns, a performance which could not be replicated for equity returns for the Finnish market. Considering the ambivalent nature of the Finnish market, we provide supporting evidence to the results for U.S. market (Chen & Ludvigson, 2009; Ghattassi, 2008) comparable to the European studies (Engsted & Møller, 2010; Engsted et al., 2010; Hyde et al., 2005).

Appendix A. Derivation of r_f under Constantinides (1990) assuming log normality

The model implies moment condition such that

$$E_t\left[\beta(\Delta c_{t-1})^{-\alpha\frac{(\Delta c_t-\theta)^{-\alpha}-\beta\theta(\Delta c_t)^{-\alpha}E_{t+1}(\Delta c_{t+1}-\theta)^{-\alpha}}{(\Delta c_{t-1}-\theta)^{-\alpha}-\beta\theta(\Delta c_{t-1})^{-\alpha}E_t(\Delta c_t-\theta)^{-\alpha}}} R_{i,t+1}\right] = 1 \qquad \text{(A.I)}$$

assuming joint log normality for all the processes in the equation, and $R_{i,t+1}$ is the return on risk free security. Then the expansion of the above equation under the law of iterated expectations and subsequent arrangements, while notating log processes in lower cases, provide the following solution

$$E(r_f) = -log\beta + \alpha(\Delta c_{t-1} + \Delta c_t + +\Delta c_{t+1}) - \frac{\alpha^2}{2}(\sigma_{c,t-1} + \sigma_{c,t} - \sigma_{c,t+1}) \qquad \text{A.II).}$$

References

Abel, A.B., 1990. Asset prices under habit formation and catching up with the Joneses, *American Economic Review Papers and Proceedings*, 80, 38-42.

Campbell, J.Y., 1999. Asset prices, Consumption, and the Business Cycle, *J.B. Taylor and M. Woodford, eds., Handbook of Macroeconomics*, 1, 1231−1303.

Campbell, J.Y., and Cochrane, J.H., 1999. By force of habit: a consumption-based explanation of aggregate stock market behaviour, *Journal of Political economy*, 107, 205−251.

Campbell, J.Y., Lo, A.W., and MacKinlay, A.C., 1997. The Econometrics of Financial Markets, USA: Princeton University Press.

Campbell, J.Y., and Shiller, R.J., 1988. Stock Prices, Earnings and Expected Dividends, *Journal of Finance*, 43(2),661-676.

Chen, X. and Ludvigson, S.C., 2009. Land of addicts? An empirical investigation of habit-based asset pricing models, *Journal of Applied Econometrics*,24(7),1057-93.

Chen, N.F., Roll, R. and Ross, S., 1986. Economic forces and the stock market, *Journal of Business*, 59(3), 383–403.

Cumby, R.E., 1990. Consumption risk and international equity returns: Some empirical evidence, *Journal of International Money and Finance*, 9, 182–192

Cochrane, J.H., 2000. *Asset Pricing*, Princeton, NJ: Princeton University Press.

Cochrane, J.H., 1997. Where is market going? Uncertain facts and novel theories, *Economic Perspectives*, 21(6), 3-37.

Cochrane, J.H., and Hansen, L.P, 1992. Asset pricing explorations for macroeconomics, *NBER Macroeconomics Annual* (MIT Press, MA).

Constantinides, G.M., 1990. Habit Formation: A Resolution of the Equity Premium Puzzle, *Journal of Political Economy* 98 (3): 519–43.

Engsted, T. and Møller, S.V., 2010. An iterated GMM procedure for estimating the Campbell-Cochrane habit formation model, with an application to Danish stock and bond returns, *International Journal of Finance and Economics*, 15(3), 213-227.

Engsted, T., Hyde, S., Møller, S.V., 2010. Habit formation, surplus consumption, and return predictability: international evidence. *Journal of International Money and Finance*, 29, 1237-1255.

Engsted, T., 1998. Evaluating the consumption-capital asset pricing model using Hansen-Jagannathan Bounds: Evidence from UK, *International Journal of Finance and Economics*, 3(4), 291-302.

Epstein L.G., Zin S.E., 1989. Substitution, risk aversion and the temporal behavior of consumption and asset returns: a theoretical framework. *Econometrica*, 57, 937–969.

Epstein, L.G., and Zin, S.E., 1991. Substitution, risk aversion, and the temporal behavior of consumption and asset returns: an empirical analysis, *Journal of Political Economy*, 99, 263−286.

Fama, E.F. and French, K.R., 1993. Common risk factors in the returns on stocks and bonds, *Journal of Financial Economics*, 33(1), 3–56.

Garcia R, Renault E´, Semenov A. 2005. A consumption CAPM with a reference level. *Working Paper, CIRANO and Universite´ de Montre´ al*.

Ghattassi, I., 2008. On the predictive power of the surplus consumption ratio, *Finance Research Letters*, 5 (1), 21-31.

Koopmans, T.C., 1960. Stationary Ordinal Utility and Impatience, *Econometrica* , 28, 287-309.

Kreps, D.M., and Porteus, E.L., 1978. Temporal Resolution of Uncertainty and Dynamic Choice Theory, *Econometrica*, 46, 185-200.

Hamori, S., 1992. Test of C-CAPM for Japan: 1980–1988. *Economics Letters*, 38, 67–72.

Hansen, L.P., 1982. Large sample properties of generalized method of moments estimators, *Econometrica*, 50(4), 1029-1054.

Hansen L.P., Heaton J., Luttmer E., 1995. Econometric evaluation of asset pricing models, *Review of Financial Studies*, 8, 237–274.

Hansen L.P., Jagannathan R., 1997. Assessing specification errors in stochastic discount factor models, *Journal of Finance*, 52, 557–590.

Hansen, L.P., and Jagannathan, R., 1991. Implications of security market data for models of dynamic economies, *Journal of Political Economy*, 99, 225−262.

Hansen, L.P., and Singleton, K.J., 1982. Generalized instrumental variables estimation of nonlinear rational expectations models, *Econometrica* , 50,1269−1288.

Hyde, S., Sherif, M., 2005. Consumption asset pricing models: Evidence from the UK, *The Manchester School*, 73, 343-363.

Hyde, S., Cuthbertson, K., Nitzsche, D., 2005. Resuscitating the C-CAPM: Empirical evidence from France and Germany, *International Journal of Finance and Economics*, 10,337.357.

Lettau, M., and Ludvigson, S., 2005. Expected returns and expected dividend growth, *Journal of Financial Economics*, 76 (2005), pp. 583–626

Lettau, M., and Ludvigson, S., 2001a. 'Consumption, aggregate wealth, and expected stock returns, *Journal of Finance*, 56(3), 815–849.

Lettau, M., and Ludvigson, S., 2001b. Resurrecting the (C)CAPM: a cross-sectional test when risk premia are time-varying, *Journal of Political Economy*, 109, 1238–1287.

Li, Y., Zhong, M., 2005. Consumption habit and international stock returns. Journal of Banking and Finance 29, 579-601.

Lintiner, J., 1965. The Valuation of Risky Assets and The Selection of Risky Investment in Stock Portfolios and Capital Budgets, Review of Economics and Statistics, 68, 452-459.

Lucas Jr., R.E. (1978). Asset prices in an exchange economy, *Econometrica* , 46,1429−1445.

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

Mehra, R. and Prescott, E.C., 2003. The Equity Premium in Retrospect, *Handbook of the Economics of Finance* ed. by G.M Constantinides, M. Harris and R. Stulz, North Holland, Amsterdam.

Mehra, R. and Prescott, E.C., 1985. The equity premium. A puzzle? *Journal of Monetary Economics*, 15, 145–161.

Newey, W.K. and West, K.D., 1987. Hypothesis testing with efficient method of moments estimation, *International Economic Review*, 28, 777–787.

Oikarinen, E. and Kahra. H., 2002. A Consumption based explanation of equity and housing property returns, *presented at Pacific RIM real estate society Conference* ,21-23 January 2002, Christchurch , New Zealand.

Roll, R.. 1977, A Critique of the Asset Pricing Theory's Tests: Part I: On Past and Potential Testability of the Theory, *Journal of Financial Economics*, 4, 129-176.

Ryder, H. E., Jr., and Heal, C. M., 1973. Optimum Growth with Intertemporally Dependent Preferences." *Review of Econmic Studies, 40,* 1- 33.

Sharpe, W.F., 1964. Capital Asset Prices: A theory of Market Equilibrium under Conditions of Risk, Journal of Finance, 19,425-442.

Viitanen, J., 2004. Consumption based CAPM model with Finnish evidence: Introductory Essay with Summaries of Other Chapters, University of Joensuu.

Virk, N.S., 2012a. Stock returns and macro risks: Evidence from Finland, *Research in International Business and Finance*, 26(1), 47-66.

Virk, N.S., 2012b. Equity Premium Puzzle: A Finnish Review, *International Journal of Economics and Finance*, 4(2), 44-55.

Weil, P., 1989. The equity premium puzzle and the risk-free rate puzzle, *Journal of Monetary Economy*, 24, 401−421.