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UNCONDITIONAL INTERNATIONAL ASSET PRICING MODELS: EMPIRICAL TESTS*

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Single and multifactor unconditional international asset pricing models are tested for the real and excess returns on Finnish size and industry portfolios using the traditional alpha intercept tests. The results support the efficiency of the global equity market portfolio, although the explanative power of the model remains low. The results also give evidence for the relevance of the global interest rate and Fama-French notion of value premium risk factors. The "pure" local market risk is also able to explain a large part of the asset returns but it does not drive out the global market riskfactor. This suggests that a segmented asset pricing model could be more appropriate for the pricing of Finnish stocks. (JEL: F30, G12, G15)

1. Introduction

A number of papers have studied international asset pricing models (see, e.g., Cho, Eun, and Senbet (1986), Korajczyk and Viallet (1989), Cumby and Glen (1990), Harvey and Zhou (1993), Dumas (1994), Bekaert and Harvey (1995), Harvey (1995), Dumas and Solnik (1995), De Santis and Gérard (1997), Fama and French (1998), Heston, Rouwenhorst, and Wessels (1999)).[1] Given the prevailing development towards more closely integrated capital markets, the results usually give at least partial support for the international asset pricing models. The results also show that multifactor asset pricing models seem to outperform the single-index capital asset pricing pricing model.

Most of the previous empirical work, however, has been conducted using country level data from the US or other highly integrated stock markets.[2] Less attention has been given to small but developed markets. Therefore, it is of interest to study international asset pricing models also using other stock markets and disaggregated portfolio data. This paper tests international asset pricing models using recent Finnish size and industry portfolio data. We begin our analysis with a standard unconditional

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[1] *See Adler and Dumas (1983) for a good review. A more recent review can be found, e.g., in Stulz (1995).*

[2] *Exceptions include, e.g., Buckberg (1995) and Harvey (1995). They both find partial support for the conditional ICAPM using data from emerging markets.*

international CAPM model and extend it to multifactor asset pricing models for real and excess returns. In particular, we are interested to study whether international sources of risks are priced in the Finnish stock market.

The Finnish stock market has experienced a recent transformation from an isolated market to an internationally accessible market which is closely integrated to the global capital markets. This development started in the early 1980s when a few domestically listed Finnish firms became listed on foreign stock exchanges. About the same time foreign investors were allowed to buy Finnish publicly listed stocks up to 20 percent of the equity capital. A few years later in the mid-1980s Finnish investors were allowed to buy foreign stocks.[3] The final step in this process took place in 1993 when all restrictions on foreign ownership were abolished (see Vaihekoski, 1997, for more details). In addition, the economical and political integration process within the EU and its monetary system has also made Finland both financially and economically more integrated with the global economy. Overall, we believe that all this makes the Finnish stock market an interesting one for tests of the international asset pricing models.

There are a few previous studies using Finnish data to study international dimensions of asset pricing and market integration. Using data for the period January 1984 to June 1985, Hietala (1989) finds the Finnish market to be segmented and that the foreign investors are willing to pay a premium for the Finnish shares, because of the possibility to diversify globally. Martikainen, Virtanen, and Yli-Olli (1993) and Bos, Fetherston, Martikainen, and Perttunen (1995) find Finnish stocks to exhibit low integration to the Swedish and especially to the US stock markets using data for 1980–1986 and 1983–1989, respectively. Harvey (1995) includes Finland in his sample and finds Finland to be one of the least correlated with the world equity market using data from 1988 to 1992. Nummelin and Vaihekoski (2000), on the other hand, studied a multifactor international asset pricing model for segmented markets using data from 1987–1996 and found evidence that global risk factors do have an important role explaining Finnish asset returns.

The rest of the paper is organized as follows. Section 2 first presents the theoretical background of international asset pricing models and their testable implications. After that the empirical testing of the models and a few econometric questions are discussed. Section 3 presents the data and its properties. Section 4 presents the empirical results and a few diagnostic tests. Finally, section 5 concludes and gives suggestions for further research.

2. Research methodology

2.1 Theoretical background

The international capital asset pricing model was originally studied by Solnik (1974). Important contributions to the theory have been made by Black (1974), Grauer, Litzenberger, and Stehle (1976), Sercu (1980), Stulz (1981), and Adler and Dumas (1983), among others. The arbitrage pricing theory (APT) was extended to the international setting originally by Ross and Walsh (1983) and Solnik (1983). The IAPT model was further extended by Levine (1989) and Ingersoll (1984), among others.[4]

The international CAPM is based on the famous CAPM model with a few additional assumptions that arise from the fact that investors are concerned about the real return of the financial assets per unit of risk rather than the nominal return in the currency of their country. First, we need to assume that the international financial market is perfect and frictionless (i.e., no barriers exist). Second, we assume the joint lognormality of the asset prices and exchange rates. Third, the asset in country j with an riskfree nominal return in currency j has a beta equal to zero. Fourth, the rate of growth of the price of the consumption good in currency j is

[3] *Of course, the economic (and financial) integration can be said to have begun much earlier when Finnish companies engaged in foreign trade and established foreign facilities.*

[4] *Errurza and Losq (1985, 1989), Basak (1996), and Bergström, Rydqvist, and Sellin (1993), among others, developed a version of the international asset pricing models under segmented market conditions.*

uncorrelated with nominal asset returns in that currency. Finally, we assume that a riskless asset exists.[5] Now, the international CAPM implies the following relation for the nominal excess returns

$$(1) \quad E[r_i] = \beta_i E[r_m]$$

where $\beta_i = Cov(r_i, r_m)/Var(r_m)$ and $E[r_i]$ and $E[r_m]$ are expected returns on asset i and the global market portfolio in excess of the riskless rate of return r_f. The global market portfolio comprises all securities in the world in proportion to their capitalization relative to world wealth (see Stulz, 1995).

However, the rate of return on default-free asset is not strictly riskless in real terms.[6] In this case we can use Black's (1972) zero-beta version of the international CAPM. It implies the following relationship for real returns (i.e., nominal returns in excess of the inflation rate)

$$(2) \quad E[R_i] = \lambda_0 + \beta_i (E[R_m] - \lambda_0)$$

where λ_0 is the expected return on the zero-beta portfolio. This is the expected return of any security that is uncorrelated with the market return.

Merton (1973) suggests that we should allow for a stochastic investment opportunity set. This gives us a theoretical basis for the multifactor intertemporal CAPM (ICAPM) in the international pricing framework.[7] Although the model does not tell us what the relevant risks are, a common approach is to consider movements in interest rates as an additional source of risk which adds yet an additional intertemporal hedge asset factor to the pricing relation.

In addition, several researchers have found that traditional pricing models are often empirically unable to explain the performance of smaller companies. Several studies have explained this as evidence in favor of the size or value premium (see e.g. Fama and French (1995) and in Europe Heston, Rouwenhorst, and Wessels).[8] Fama and French (1998) compare the standard international CAPM model against a model where they have added a global value premium factor and find the additional value premium factor highly relevant. We also test for this possibility by adding a value premium factor as the third source of risk in our international multifactor asset pricing model (APM).

Now we can write the unconditional pricing model as follows

$$(3) \quad E[R_i] = \lambda_0 + \sum_{k=1}^{K} \beta_{ik} \lambda_k$$

where K is the number of risk factors, β_{ik} is asset i's sensitivity to the kth risk factor, and λ_k is the expected risk premium on kth risk factor in excess of λ_0. The λ_0 is the expected *real* return of any security that is uncorrelated with each of the K risk factors in the model. If a riskless asset exists, then λ_0 is the rate of return of this asset.

2.2 Pricing Models and Testable Implications

There are several approaches to testing (international) asset pricing models.[9] Here we concentrate on testing the time-series implications of the unconditional models. This can be motivated by our interest in the ability of the model to explain the cross-sectional differences in asset returns.

We start with the traditional testing procedure where we add an alpha parameter to the pricing equation and test the implied restrictions on this parameter. Pricing models for excess returns imply that under the null alpha is equal to zero. Correspondingly, the zero-beta pric-

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[5] I.e., we assume that investors can freely lend and borrow in units of the numeraire consumption good (see Stulz, 1995).

[6] In empirical studies the riskless asset is usually proxied by the rate of return on Treasury Bills which is subject to inflation risk.

[7] Alternatively, a multifactor model could be justified by the APT theory.

[8] Although size and value premiums are somewhat different, Fama and French (1995) argue that their value premium factor proxies for relative distress of the firms similar to the size premium.

[9] The most commonly used tests of asset pricing models are implementations either of the cross-sectional test, originally by Fama and MacBeth (1973), the zero-intercept test of time-series regression, originally by Gibbons (1982), or the pricing kernel approach.

ing model implies the following restriction $\boldsymbol{\alpha} = (\boldsymbol{\iota}-\boldsymbol{\beta})\lambda_0$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $(N \times 1)$ vectors, and $\boldsymbol{\iota}$ is a $(N \times 1)$ vector of ones.[10] Rejecting the null hypothesis rejects the (multi)factor portfolio mean-variance efficiency and the model.

A particular pricing model also implies the relevant sources of risk. Hence, we are interested to test whether the implied risk factors are statistically relevant to the pricing of the assets. Pricing models also implicitly imply that these risk factors are the only priced risk factors and no other risk factors should appear to be significant in the pricing equation. In particular, international asset pricing models imply that local risk factors should not be priced.

2.3 Testing Procedure and Econometric Questions

We evaluate the traditional unconditional pricing models using Ordinary Least Squares (OLS), Maximum Likelihood (ML), and the GMM metric. First we use the OLS estimation to get the alpha and beta parameter estimates and the traditional overall model performance measure (adjusted R-square). Since previous studies have reported evidence that Finnish return series exhibit autocorrelation, heteroskedasticity, and nonnormality (see, e.g., Vaihekoski, 1998), we use the Newey and West (1987) autocovariance and heteroskedasticity consistent covariance matrix estimator to adjust the standard errors respectively. Autocorrelation adjustment can also be motivated if there are reasons to believe that the thin trading effect is evident in the data as is the case in the Finnish market.

Using the covariance matrix estimator from the OLS estimation, we can also calculate a multivariate Wald test statistic to test the null hypothesis that a certain parameter is cross-sectionally zero. The Wald test statistic is calculated as follows

$$(4) \quad W = (\mathbf{R}\hat{\mathbf{b}} - \mathbf{r})'(\mathbf{R}\hat{\mathbf{C}}\mathbf{R}')^{-1}(\mathbf{R}\hat{\mathbf{b}} - \mathbf{r}) \sim \chi^2_N$$

where $\mathbf{R}$ and $\mathbf{r}$ are $(N \times M)$ and $(N \times 1)$ restriction matrices, $\hat{\mathbf{b}}$ is a $(M \times 1)$ vector of coefficient estimates, $\hat{\mathbf{C}}$ is a $(M \times M)$ Newey-West covariance matrix estimator, N is the number of test assets (or restrictions), and M is the number of equations (assets) times the number of parameters estimated for each equation (alpha and betas). In the tests, $\mathbf{R}$ is a matrix of zeros, except for those parameters that we want to pick for the test and $\mathbf{r}$ is a vector of zeros (due to null hypothesis).[11]

In contrast, we cannot use the OLS framework to estimate the pricing model for the real returns. The ML framework, however, can be easily used to estimate the model. Using the approach in Campbell, Lo, and MacKinlay (1997), we calculate seven test statistics for our pricing model. The first four test statistics J_0, J_1, J_2, and J_3 are different variations of the test for the excess return formulation of the asset pricing model. The first test statistic J_0 is a Wald test defined as follows

$$(5) \quad J_0 = T\left[1 + \frac{\hat{\mu}_m^2}{\hat{\sigma}_m^2}\right]^{-1} \hat{\boldsymbol{\alpha}}'\boldsymbol{\Sigma}^{-1}\hat{\boldsymbol{\alpha}}$$

where T is the number of time series observations, $\hat{\mu}_m$ is the sample mean excess market return, $\hat{\sigma}_m^2$ is the sample variance of the market excess return, $\hat{\alpha}$ is a vector of estimated alpha parameters, and $\boldsymbol{\Sigma}$ is the return variance-covariance matrix. Under the null hypothesis $\alpha = 0$, J_0 is distributed as a chi-square distribution with N degrees of freedom. The next test statistic J_1 is similar to J_0, but instead of T, the latter part of the equation is multiplied with $(T-N-1)/N$ to get a finite-sample F-test. J_1 is distributed central F with N degrees of freedom in the numerator and $(T-N-1)$ degrees of freedom in the denominator.

The next test statistics J_2 is a likelihood ratio test and defined as follows

$$(6) \quad J_2 = T\,[\log|\hat{\boldsymbol{\Sigma}}^*| - \log|\hat{\boldsymbol{\Sigma}}|]$$

where $\hat{\boldsymbol{\Sigma}}^*$ and $\hat{\boldsymbol{\Sigma}}$ are the maximum likelihood estimators of the residual covariance matrix for the unconstrained and constrained models, re-

[10] *Note that the alphas and betas are not the same for excess and zero-beta models.*

[11] *For more information see Greene (1997).*

spectively. J_2 is distributed chi-square with N degrees of freedom. Similar to J_1, we can improve the finite-sample properties of the J_2 by using $(T–N/2–2)$ instead of T in (6) to get a modified test statistic J_3.

The next three test statistics, J_4, J_5, and J_6, are test statistics for the zero-beta pricing model using real returns. J_4 and J_5 are similar to J_2 and J_3, but now we have one more constant to be estimated (expected zero-beta rate) so we lose one degree of freedom. J_6 is similar to J_1 and distributed central F with N degrees of freedom in the numerator and $(T–N–1)$ degrees of freedom in the denominator. These test statistics also extend to multifactor models, but we lose $K–1$ degrees of freedom.[12]

The ML tests above require that asset returns are i.i.d. and multivariate normal conditional on the factor portfolio returns. If we want to relax these assumptions we can construct more robust tests in the GMM framework. The GMM has several advantages that have made it popular in financial model analysis. First, joint estimation of all model parameters reduces the errors-in-variables problem. Second, the GMM does not rely upon the assumption of normally distributed asset returns; the disturbance term can be both serially dependent and conditionally heteroskedastic. In fact, the only requirements are that the data is strictly stationary and ergodic. (Hansen, 1982; Ferson, Foerster and Keim, 1993; MacKinlay and Richardson, 1991).

To test corresponding zero-intercept restrictions set by the pricing models in the GMM framework, we write the following error terms for assets $i=1,..,N$.

$$(7) \quad u_{it}=R_i-\alpha_i-\boldsymbol{\beta}'\boldsymbol{\lambda}_K$$

where the model implies the following moment condition $E[u_{it}|1,\boldsymbol{\lambda}_K]=0$. If a riskless asset exists, returns and risk factor returns are in excess of the riskless rate of return. In a general case we can write the moment conditions as a function of the parameters, or $E[f_t(\boldsymbol{\theta})|1,\boldsymbol{\lambda}_K]=0$. Using this notation, the moment condition can be written as the sample average

$$(8) \quad \mathrm{g}_T(\boldsymbol{\theta})=\frac{1}{T}\sum_{t=1}^{T}\mathrm{f}_t(\boldsymbol{\theta})$$

where the GMM estimator is chosen to minimize the quadratic form

$$(9) \quad \mathrm{g}_T(\boldsymbol{\theta})'\mathrm{W}\mathrm{g}_T(\boldsymbol{\theta})$$

where **W** is a positive definite weighting matrix.

In equation (7) we have $(N + N \times K)$ parameters to be estimated and equally many moment conditions. The system is thus exactly identified and the GMM estimator will not depend on **W**. In fact, the parameter estimates are similar to the OLS estimates. However, their standard deviations differ from the OLS and they are affected by the choice of the weighting matrix. Again, we employ the Newey-West (1987) weighting matrix in the estimation to correct the standard deviations for the effects of autocorrelation and heteroskedasticity.

To test the model we insert the appropriate restrictions implied by the pricing model on the α_i term. Now the system becomes overidentified and we can test the restriction using the LR test statistic which is distributed under the null as χ^2_N where the number of restrictions equals N.[13]

The implementation of the overidentified GMM includes, however, a few practical difficulties. One of the main difficulties is the optimization problem. The dimensionality of the estimation system can become too large to be estimated reliably using the GMM. And even when the solution is found numerically, the solution may not converge to the global minimum or converge at all. (Zhou, 1994). However, in our case we have only a small number of parameters to estimate. We also try to mitigate this problem by testing the solution to iteration with several different initial parameter values. In general, we find our estimates to be stable and convergence is easily atteined.

Another problem is caused by the somewhat limited knowledge of the small-sample properties of the GMM estimator. We consider later

[12] *See Campbell, Lo, and MacKinlay (1997) for more information.*

[13] *Note that when the LR test is conducted, we have to use the same weighting matrix as in the original estimation.*

the small-sample behavior of our GMM test statistics.

3. Data

The estimation period covers 144 months of data from January 1987 to December 1998. Competitively determined Finnish short-term interest rates have existed during the whole time period. In particular, a proxy for the representative agents' nominally risk-free return is taken to be the 1-month interbank money market rate – the Helsinki Interbank Offered Rate (Helibor) – on the last trading day of each month *t*–1.[14] Real returns are calculated by substracting the monthly inflation rate from the nominal returns. The inflation rate series is taken from the ETLA (The Research Institute of the Finnish Economy) database. We employ continuously compounded asset returns throughout the paper.

3.1 Global risk factors

We employ up to three risk factors in our international asset pricing model to represent economic risks. As discussed in Section 2, these risk factors have been chosen based on theoretical and empirical considerations. First, we employ a global market portfolio as represented by the monthly return ($r_{wm,t}$) on the Morgan Stanley Capital International (MSCI) World Equity Market index where gross dividend are invested back into the market.[15] Second, we consider the global bond market return ($r_{wb,t}$) calculated from the Salomon Brothers World Government Bond Index. Third, we use a proxy for the global value premium ($r_{vp,t}$) measured as the difference between return on the MSCI World Value and Growth Equity indices.

The MSCI and Salomon Brothers indices used here are originally expressed in dollar terms. Since we take the perspective of a Finnish investor and want to express all returns in Finnish markkas (FIM), we converted them into FIM with the end-of-the-month FIM/USD exchange rate. The MSCI index is calculated at 4:00 pm, US eastern standard time (GTM-0600).[16] The currency index is calculated every day at noon Finnish time (GTM+0200). This produces a slight timing difference but its magnitude is expected to be small in the monthly data.

As seen in Table 1, the mean annual return on the world equity portfolio has been more than 12 percent, while the mean nominal annual return in FIM for the world bond markets has been 9.4 percent. The value rate risk factor shows an annual premium of 0.8 percent. During the sample period the local riskfree interest rate has been high, on average, 8.5 percent, implying a global market risk premium of four percent. The inflation has stayed quite low during the sample period, on average, 2.9 percent. There is evidence of significant first-order autocorrelation in the equity related factors and we reject normality for all risk factors using the Bera-Jarque test statistic.

3.2 Test assets

In this paper, we study the monthly value-weighted returns on six portfolios sorted by size (total equity market capitalization) and seven industry portfolios.[17] Stock returns series are adjusted for stock splits and share issues and dividends are accounted for by assuming that dividends received on a stock are reinvested in the same stock. Portfolios consists of all stocks listed on the Helsinki Stock Exchange during December 1986 – December 1998. For each firm only one series – the most actively traded – was chosen. Thereby, we use some future information but only for choosing which one of the usually multiple stock series to employ for each company.[18]

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[14] Helibor rates are used as a proxy for riskless rate of returns since trading on T-bills started later in the 1990s. Moreover, Finnish government backs up banks' CDs so they have been widely considered as virtually risk-free.

[15] The index is not corrected for cross-country holdings but it will probably have only a negligible effect on our results (see e.g., Bansal, Hsieh, and Viswanathan, 1993).

[16] A possible exception to the timing difference is the daylight saving time adjustment, but during most years it should have affected all timing similarly.

[17] The number of portfolios is limited since the number of companies available varied from 50 to 83 (mean 66) during the sample period.

[18] More information of the portfolio construction can be found from Vaihekoski (2000).

Table 1. Descriptive statistics of the risk factors.

The descriptive statistics are calculated for risk factor variables. The *p*-value for the Bera-Jarque test statistic of the null hypothesis of normal distribution is provided in the table. World stock market returns are calculated from the Morgan Stanley Capital International (MSCI) World Equity Index. World bond returns are calculated from the Salomon Brothers World Government Bond Index. Both variables are converted into Finnish currency. The world value premium is the different in returns on MSCI World Growth and Value indices. The mean and standard deviation of the risk variables are annualized in the table. The sample size is 144 monthly observations from 1987 to 1998.

Time Series	Symbol	Mean	Standard Deviation	Skewness	Excess kurtosis	Normality (p-value)	Autocorrelation[a] ρ_1	ρ_2	ρ_3	ρ_{12}	Q(12)[b]
Economic risk variables (annualized)											
Return on world stock index	$R_{wm,t}$	0.124	0.173	0.965	2.868	<0.001	0.225*	–0.008	–0.098	–0.052	0.066
Return on world bond index	$R_{wb,t}$	0.094	0.093	1.703	7.429	<0.001	0.157	–0.085	0.028	–0.074	0.002
World value premium	$R_{vp,t}$	0.008	0.056	–0.217	1.347	0.004	0.160	0.051	0.023	0.118	0.130
Risk-free return	$R_{f,t}$	0.085	0.012	0.211	–1.233	0.054	0.969*	0.945*	0.921*	0.816*	<0.001
Inflation rate	I_t	0.029	0.012	1.092	1.417	<0.001	0.245*	0.051	0.296*	0.566*	<0.001
Correlation matrix of the risk factors		$R_{wm,t}$	$R_{wb,t}$	$R_{vp,t}$	$R_{f,t}$	I_t					
Return on world stock index	$R_{wm,t}$	1.000									
Return on world bond index	$R_{wb,t}$	0.515	1.000								
World value premium	$R_{vp,t}$	–0.114	–0.020	1.000							
Risk-free return	$R_{f,t}$	–0.031	0.120	0.014	1.000						
Inflation rate	I_t	–0.058	–0.057	0.075	0.396	1.000					

[a] Standard error for autocorrelation coefficients with lag *q* is given by $\sqrt{(1+2\rho_1^2+\ldots+2\rho_q^2)}/\sqrt{T}$. Values significantly different from zero are marked with *.
[b] The *p*-value for the Ljung and Box (1978) test statistic for the null that autocorrelation coefficients up to 12 lags are zero is provided.

When forming the size portfolios, we have tracked listings and delistings on a monthly basis and our portfolios are rebalanced each month t based on market values at the end of month *t–1*. Hence, a given company's shares only have to be listed for two consecutive months to qualify for our sample. The last (incomplete) listing periods are excluded from the series. All size portfolios have roughly the same number of stocks at each month *t*.[19] Industry portfolios are formed by sorting firms at the end of the calendar year into groups based on their industry classification given by Talouselämä[20].

Summary statistics for the portfolio nominal return series are provided in Table 2. Mean and standard deviation are scaled by 12 and the square root of 12 to show them in annual terms. Overall, we can see that average returns on some of the portfolios is less than the riskless return, especially for smaller firms and for certain industries (banking and other financial, housing and construction). This is probably due to the recession in the Finnish economy in the early 1990s which hit those companies hardest that were more dependant on the domestic markets.

Negative average realized excess portfolio returns and near zero excess market returns raise questions regarding the validity of our results. It is true that using only ten years of data is not enough to estimate the market premium very accurately. In general, our question is whether we can proxy expected returns using realized returns. However, it is reasonable to assume that the expected excess returns have been positive for the most part even if the sample realizations have been negative. Thus, we follow the standard approach used in most of the previous studies.

The asset returns show also signs of nonnormality. The Bera-Jarque test for normality rejects the null hypothesis of normal distribution for seven of the thirteen portfolios. There is also evidence of serial correlation in the portfolio

[19] *An equal number of stocks is assigned to each portfolio except "extra" stocks (i.e., Mod(N/P) > 0, where N is the number of companies and P is the number of portfolios) are assigned sequentially, one per portfolio, beginning from the smallest portfolio.*

[20] *Talouselämä is the leading weekly business magazine in Finland. They publish yearly a roundup review of the biggest 500 Finnish companies.*

Table 2. Summary statistics for the monthly nominal portfolio returns.

Descriptive statistics are calculated for the nominal portfolio returns. The *p*-value for the Bera-Jarque test statistic of the null hypothesis of normal distribution is provided in the table. The mean and standard returns are annualized (multiplied with 12 and the square root of 12, respectively). The sample size is 144 monthly observations from 1987 to 1998.

Time series	Mean	Standard Deviation	Skewness	Excess kurtosis	Normality (p-value)	Autocorrelation[a]				
						ρ_1	ρ_2	ρ_3	ρ_{12}	Q(12)[b]
Size portfolios										
Largest	0.135	0.269	–0.321	–0.034	0.290	0.148	0.056	0.063	0.065	0.003
2	0.056	0.254	–0.303	0.829	0.042	0.269*	–0.021	0.095	0.181*	<0.001
3	0.074	0.276	–0.271	2.331	<0.001	0.281*	–0.094	0.005	0.088	0.001
4	0.082	0.240	0.269	2.035	<0.001	0.274*	0.041	0.139	0.168	0.001
5	0.027	0.225	–0.230	0.772	0.089	0.245*	–0.007	0.063	0.139	0.002
Smallest	0.026	0.253	0.670	2.474	<0.001	0.165*	0.046	0.096	0.019	0.068
Industry portfolios										
Banking & Other Financial	–0.052	0.332	0.384	1.831	<0.001	0.179*	0.039	0.162	–0.032	0.001
Forestry	0.081	0.253	–0.030	0.396	0.619	0.130	0.075	–0.155	0.117	0.002
Trade & Transport	0.052	0.235	–0.264	0.501	0.205	0.231*	–0.001	0.201*	0.231*	<0.001
Metal & Electronics	0.189	0.287	0.104	–0.114	0.845	0.094	–0.115	0.131	0.134	0.041
Food Industry	0.076	0.248	–0.255	1.786	<0.001	0.095	0.104	0.028	0.142	0.382
Housing & Construction	–0.076	0.319	–0.412	2.151	<0.001	0.216*	0.002	0.129	0.111	0.011
Multi-Business	0.128	0.316	–0.284	0.787	0.059	0.261*	0.048	0.096	0.095	<0.001

[a] Autocorrelation coefficients significantly (5%) different from zero are marked with an asterisk (*).
[b] The *p*-value for the Ljung and Box (1978) test statistic for the null that autocorrelation coefficients up to 12 lags are zero is provided.

returns. For example all but the largest size portfolios have significant first-order correlation coefficients. Ljung-Box test statistic rejects hypothesis of zero autocorrelation coefficients up to twelve lags for eleven out of thirteen portfolios. Therefore, we choose to set the number of lags for the Newey-West (1987) estimator to three.[21]

4. *Empirical results*

4.1 International CAPM

Table 3 shows the results for the one-factor international CAPM model where we have used the OLS to estimate the model using excess returns on size and industry portfolios, respectively. In general, the explanative power of the model (adjusted R-square) is rather low (between 7.6 and 31.2 percent for size portfolios and between 9.7 and 28.2 percent for industry portfolios) and on average at par with what Harvey (1995) found for the whole Finnish equity market (21.9 percent).[22]

All portfolios exhibit relatively low global market risk sensitivity. Size portfolio betas range from 0.410 for the smallest size portfolio to 0.881 for the largest size portfolio. Contrary to US studies, market betas seem to decrease for smaller companies (Heston, Rouwenhorst, and Wessels have similar results). This is partly due to the thin trading effect in the portfolio returns, but Scholes and Williams (1977) adjusted betas (not reported) still remain below one even though are they are generally higher than the betas reported in the table.[23] Our sample is probably too short to provide a good estimate of the long-run average unconditional betas using realized returns. In addition, the deep economic downturn in the Finnish economy dur-

[21] Newey-West estimator requires a bandwidth (or lag length) to represent the number of observations that receive non-zero weight at each point in the estimation. Three is less than some suggestions (see, e.g., Hamilton, 1994) of setting it to close to $T^{1/3}$ if no prior information is available. With 144 observations it would be 5.2. However, based on the autocorrelation structure of the information variables and asset returns, we set it to three.

[22] We also performed the tests for the whole Finnish market. The adjusted R-square was 30.9 percent and market beta 0.816 (compared to 0.667 of Harvey).

[23] Scholes and Williams (1977) size portfolio betas range from 0.919 (second largest size portfolio) to 0.475 (smallest portfolio).

Table 3. Tests of Unconditional International Capital Asset Pricing Model.

One-factor international capital asset pricing model is tested using excess and real returns of six size and seven industry portfolios. Excess and real returns are formed by subtracting the Helibor interest rate or *ex post* inflation from the nominal returns, respectively. In panels A and C we report results from the OLS estimation where we have used excess returns and the Newey-West (1987) autocovariance and heteroskedasticity consistent covariance matrix with three lags. World stock market returns are calculated from the Morgan Stanley Capital International (MSCI) World Equity Index. Standard errors are given below the parameter estimates. In panels B and D we report various ML and GMM test statistics for the restrictions set by the pricing model on excess and real returns. The sample size is 144 observations from January 1987 to December 1998.

Panel A: Size Portfolios	Average		Beta	Adj. R^2
	Excess return	Pricing error α	Market risk	
Largest	0.004	0.001	0.881*	0.312
		0.007	0.101	
2	–0.002	–0.005	0.770*	0.269
		0.006	0.107	
3	–0.001	–0.003	0.687*	0.179
		0.007	0.121	
4	–0.000	–0.002	0.518*	0.132
		0.006	0.092	
5	–0.005	–0.006	0.463*	0.124
		0.006	0.082	
Smallest	–0.005	–0.006	0.410*	0.076
		0.006	0.107	
Wald-test coefficient[a]		3.230	>99.999	
(p-value)		0.780	<0.001	

Panel B: Tests	Excess returns	(p-value)		Real returns	(p-value)
J0	5.020	0.541	J4	3.381	0.641
J1	0.796	0.575	J5	3.263	0.659
J2	4.935	0.552	J6	0.588	0.739
J3	4.763	0.575			
GMM-test (LR)	4.566	0.601		3.770	0.583

[a] Wald cross-sectional test with the Newey-West (1987) heteroskedasticity and autocorrelation consistent matrix estimator with three lags.

ing the early 1990s is reflected in the results.[24] The lowest global risk beta is 0.538 for the industry portfolios and it is shared by the two *a priori* most domestic industries: Food industry and Housing & Construction industry. On the other hand Banking & Other Financial industry surprisingly has the highest global market risk sensitivity: 0.831.

In panels B and D, we report the results from the tests on the alpha-parameter restriction using both excess and real returns on size and industry portfolios, respectively. Somewhat surprisingly, the results for the size portfolios do not reject the mean-variance efficiency of the global market portfolio at the standard five percent level using either real or excess returns (Wald test on alpha, J_0–J_6 and GMM test statistics have p-values higher than 5 percent). The results thus present evidence in favor of the international CAPM.

However, results for the industry portfolios are not that favorable for the international CAPM. All test statistics either reject the model or are close to rejecting it. In addition, comparing average excess returns and alpha (intercept) estimates, we can see that the absolute

[24] *One alternative explanation is also that the monthly updating of the portfolio contents lowers betas for the smaller portfolios when compared to yearly updated portfolios. We studied this using both types of portfolios and found out that this is not the case.*

Table 3. *Continued*

Panel C: Industry Portfolios	Average		Beta	Adj. R^2
	Excess return	Pricing error α	Market risk	
Banking & Other Financial	–0.011	–0.014	0.831*	0.191
		0.008	0.182	
Forestry	–0.000	–0.003	0.782*	0.282
		0.005	0.107	
Trade & Transport	–0.003	–0.005	0.600*	0.189
		0.006	0.111	
Metal & Electronics	0.006	0.006	0.828*	0.248
		0.006	0.138	
Food Industry	–0.001	–0.002	0.538*	0.135
		0.006	0.128	
Housing & Construction	–0.013	–0.015	0.538*	0.097
		0.009	0.128	
Multi-Business	0.004	0.001	0.783*	0.178
		0.008	0.124	
Wald-test on coefficients[a]		7.709	>99.999	
(p-value)		0.359	<0.001	

Panel D: Tests	Excess returns	(p-value)		Real returns	(p-value)
J0	14.378	0.045	J4	11.051	0.087
J1	1.940	0.068	J5	10.629	0.101
J2	13.704	0.057	J6	1.470	0.183
J3	13.181	0.068			
GMM-test (LR)	12.144	0.096		10.001	0.125

[a] Wald cross-sectional test with the Newey-West (1987) heteroskedasticity and autocorrelation consistent matrix estimator with three lags.

pricing errors are typically larger than absolute average excess returns. In fact, the avarage absolute pricing error for the industry portfolio is 54 basis points compared to that of 28 for the size portfolios. Housing & Construction industry portfolio also has the largest pricing error 1.5% per month. Since alpha is generally denoted as the pricing error (or Jensen's alpha), this together with relatively low *R*-squares cast doubt on the model's ability to price all assets adequately.

Comparison of the results across test statistics reveals that their finite sample inferences are somewhat different. Since we know the exact finite sample distribution of J_1 under the null, we keep that as the basis of our comparison. The J_2 test statistic is higher than J_1 (i.e. p-value is lower) in all cases indicating that the asymptotic likelihood ratio test tends to reject too often. The finite-sample adjustment to J_2 works well as J_3 is almost identical to J_1. Robust estimation with the GMM shows surprisingly different results for the real and excess returns. For real returns the p-values drop, but for the excess returns they are higher. Therefore, we cannot draw conclusions whether the GMM test rejects the model more often.

The surprising inability of the test statistics to reject the mean-variance efficiency of the global equity market could be partly caused by the low power of the tests. Campbell, Lo, and MacKinlay (1997) provide some results regarding the power of these tests using the J_1 test statistic. They compare the null hypothesis against different hypotheses of the real average market portfolio return. When $T=120$ and $N=5$, they find the power of the J_1 statistic to be quite low, it varies between 0.106–0.332 for different alternative assumptions.

Table 4. Tests of Unconditional Three-Factor IAPM Model

Three-factor international asset pricing model is tested using excess and real returns of six size and seven industry portfolios. Excess and real returns are formed by subtracting the Helibor interest rate or *ex post* inflation from the nominal returns, respectively. In panels A and C we report results from the OLS estimation where we have used excess returns, a global three-factor pricing model with global equity market, interest rate and value risk factors, and the Newey-West (1987) autocovariance and heteroskedasticity consistent covariance matrix with three lags. Standard errors are given below the parameter estimates. In panels B and D we report various ML and GMM test statistics for the restrictions set by the pricing model on excess and real returns. The sample size is 144 observations from January 1987 to December 1998.

Time series	Average		Beta			Adj. R^2		
	Excess return	Pricing error α	Market risk	Interest rate risk	Value premium	One-factor model	Two-factor model	Three-factor model
Panel A: Size Portfolios								
Largest	0.004	0.006	1.071*	–0.599*	0.819*	0.312	0.336	0.361
		0.006	0.135	0.255	0.321			
2	–0.002	0.002	0.980*	–0.631*	1.170*	0.269	0.299	0.363
		0.005	0.151	0.263	0.315			
3	–0.001	0.004	0.817*	–0.345	1.152*	0.179	0.189	0.232
		0.005	0.146	0.319	0.352			
4	–0.000	0.002	0.761*	–0.811*	0.604	0.132	0.193	0.208
		0.006	0.115	0.268	0.322			
5	–0.005	–0.003	0.637*	–0.564*	0.612*	0.124	0.154	0.172
		0.005	0.113	0.216	0.309			
Smallest	–0.005	–0.001	0.627*	–0.699*	0.795*	0.076	0.115	0.140
		0.007	0.124	0.254	0.328			
Wald-test		2.432	>99.999	35.914	44.330			
(p-value)		0.876	<0.001	<0.001	<0.001			

Panel B: Tests	Excess returns	(p-value)		Real returns	(p-value)
J1	0.664	0.675	J4	3.424	0.635
J2	4.223	0.647	J5	3.258	0.660
J3	4.018	0.674	J6	0.608	0.724
GMM-test (LR)	5.316	0.504	GMM-test (LR)	1.803	0.876

4.2 International multifactor APMs

Table 4 shows the results from the tests of the international multi-factor APM models. In general, adding global interest rate risk and value premium factors to the pricing model seems to increase the overall explanative power somewhat (*R*-squares increase typically about 3–5 percentage points). Overall all portfolios show consistent sensitivity to these two global risk factors having negative interest rate sensitivity and positive value premium sensitivity.

Using a multivariate Wald test statistic (4) to test the joint significance of the risk-factor loadings, we find both of these additional risk-factors appear significantly in the pricing model of the size portfolios. The value premium factor is found to be surprisingly strong in the model – all but the fourth largest portfolio have significant exposure to the international value factor.

Industry portfolios also show evidence in favor of the additional risk factors, although only one portfolio shows significant exposure to the interest rate risk. Industry portfolios also show much more dispersion in their risk sensitivity to these additional sources of risk. Analyzing individual portfolios we can see that somewhat surprisingly the banking industry does not show significant interest rate risk sensitivity. This could be caused by the deep banking sector crisis in Finland the early 1990s. Interestingly – albeit expectably – we find the Housing & Con-

Table 4 *continued*

Time series	Average		Beta			Adj. R^2		
	Excess return	Pricing error α_i	Market risk	Interest rate risk	Value premium	One-factor model	Two-factor model	Three-factor model
Panel C: Industry Portfolios								
Banking & Other Financial	–0.011	–0.007	1.027*	–0.580	1.189*	0.191	0.202	0.237
		0.008	0.258	0.335	0.390			
Forestry	–0.000	0.001	1.901*	–0.371	0.558	0.282	0.289	0.300
		0.005	0.114	0.328	0.306			
Trade & Transport	–0.003	0.002	0.765*	–0.481	1.040*	0.189	0.206	0.264
		0.005	0.142	0.295	0.333			
Metal & Electronics	0.006	0.010	1.037*	–0.684	0.686*	0.248	0.276	0.289
		0.006	0.230	0.363	0.345			
Food Industry	–0.001	0.003	0.619*	–0.203	0.826*	0.135	0.132	0.162
		0.005	0.144	0.232	0.338			
Housing & Construction	–0.013	–0.007	0.923*	–1.238*	1.387*	0.097	0.177	0.231
		0.008	0.203	0.319	0.489			
Multi-Business	0.004	0.009	0.963*	–0.510	1.322*	0.178	0.187	0.238
		0.007	0.133	0.336	0.388			
Wald-test		6.522	>99.999	28.596	51.964			
(p-value)		0.480	<0.001	<0.001	<0.001			

Panel D: Tests	Excess returns	(p-value)		Real returns	(p-value)
J1	1.429	0.199	J4	5.172	0.522
J2	10.364	0.169	J5	4.902	0.556
J3	9.824	0.199	J6	0.788	0.598
GMM-test (LR)	10.998	0.139	GMM-test (LR)	6.667	0.353

struction industry to show the highest interest rate sensitivity due to the natural dependency of its business on interest rates. Somewhat surprisingly the Housing & Construction industry exhibits also the highest significant value sensitivity.

In panels B and D of Table 4, we report the results for the alpha-parameter restriction tests using both excess and real returns of size and industry portfolios. We cannot reject the multifactor efficiency of the risk factors in using any of the test statistics. This could be caused by the low power of the test statistics. Comparing results for the excess and real returns show again that the model is more likely to be rejected when tested using excess returns. This can indicate some kind of problems with our proxy for the riskless interest rate.

4.3 Diagnostic tests

4.3.1 Small sample properties of alpha restriction tests

As stated earlier, the short sample size raises some questions of the small sample properties of the test statistics. Although we can see some guidance from the results by Campbell, Lo, and MacKinlay (1997), these results are not fully suitable for judging the behavior of the test statistics in the Finnish market. Namely, size and power tests in Campbell, Lo, and MacKinlay (1997) were conducted using the ML estimator under the assumption of normally distributed and independent observations. However, central elements of stock price behavior in many small stock markets are autocorrelation, heteroskedasticity, and nonnormality in the return series. There are several reasons for these, the most notable being the nonfrequent trading (thin trading).

Therefore, we conduct a simulation experiment to examine the small sample size properties of the maximum likelihood J_1 test statistic and three GMM test statistics, namely the likelihood ratio test statistic, Wald test statistic (18), and J_7 test statistic in Campbell, Lo, and MacKinlay (1997). We use simulated samples for the excess benchmark (market portfolio) and size portfolio returns. Generated samples have 144 observations (the same as the sample size in this study). In the GMM estimations, a HAC covariance matrix estimator is used with three lags as in this paper.

To generate simulated data series we follow a commonly used approach and fit a vector autoregressive process of the second order for the benchmark returns.[25] Then we generate new samples for the benchmark using parameter estimates and bootstrapping from the original benchmark residuals. This preserves the unconditional distributional characteristics of the original benchmark asset data.

To simulate test asset returns we first use the approach in Bekaert and Urias (1996). We model the test asset returns as a linear function of the simulated benchmark returns where parameters are from the original data and a normal, homoskedastic error term with variance equal to the original residual variance such that the CAPM/mean-variance efficiency restriction is imposed. We also use slight variation of their approach where the error term also maintains the cross-correlation structure of the original residuals. In addition, we construct a simulated asset return series following MacKinlay and Richardson (1991) where the error term is drawn with replacement from the original disturbance vector. Since drawings are done for all assets at the same time, the simulated series relies on the temporal i.i.d. assumption but maintains any contemporaneous cross-correlation and conditional heteroskedasticity in the original data.

Third, we use a method to maintain also the autocorrelation pattern in the original return series. We estimate the following model for asset returns where we allow for an autoregressive error process (AR(3) process) in order to model for the serial correlation in the asset returns.

$$(10) \quad r_{it} = b_i r_{mt} + e_{it}$$

$$(11) \quad e_{it} = \sum_{l=1}^{3} \rho_l e_{it-l} + u_{it}$$

Then we produce new test asset returns using original beta estimates and new benchmark returns. We then add the autoregressive pricing error using the parameters estimated from the original data with a normally distributed u_{it} residual term where variance is set to match that of the original data.

Finally, we follow the previous method but bootstrap u_{it} residuals from the original residual to maintain both the contemporaneous cross-correlation and the conditional heteroskedasticity in the original return series.[26] Thus, we have four different simulated series for asset returns where the mean and variance dynamics are completely driven by the benchmark returns and the mean-variance efficiency is imposed. The model and test statistics are then estimated using the same simulated data in each iteration with 1000 replications.

Results are given in Table 5. The three GMM test statistics show large deviations in their size properties. The J_7, Wald, and Likelihood Ratio tests work rather well for n.i.i.d. errors, but in other cases their size properties deteriorate. This is somewhat surprising since the GMM test statistics (J_7) should be robust with respect to deviations from the i.i.d. assumption. Surprisingly the Maximum Likelihood J_1 test statistic works well for n.i.i.d. returns (size is less than the actual 5 percent) and is relatively good in the other cases as well. In general we can conclude that our inability to reject the international CAPM model in Table 3 is not entirely caused by the limited power of the J_1 test statistic in the sample.

[25] *We set the order of the process to two on the basis of the fit.*

[26] *Note that the temporal cross-correlation structure is lost in these experiments. One alternative to the present approach would be to bootstrap sample blocks of data as in Hall and Horowitz (1996). However, this would presents some additional problems.*

Table 5. Short Sample Simulation Evidence of the Test Statistics.

The empirical size (in percent) is given for a 5 percent size test with simulated excess size portfolio and market risk factor returns.. Tests are conducted for one maximum likelihood test statistic (J_1) and three GMM test statistics (J_7, Wald, and Likelihood Ratio). They all test the alpha intercept restriction (mean-variance efficiency) in the empirical model. GMM estimation is performed with a Newey-West (1987) weighting matrix with three lags. The sample size is 144 observations. A total of 1,000 Monte Carlo replications are employed in the tests. In each replication all test statistics are calculated and saved.

	ML Test	GMM Tests		
	J_1	J_7	Wald	LR
N.i.i.d. homo-skedastic errors	2.30	12.30	9.00	7.60
N.i.i.d. errors	1.90	9.40	12.50	4.60
Bootstrapped errors	14.00	44.70	30.10	30.70
AR(3) with n.i.i.d. residual	6.20	24.20	17.90	19.90
AR(3) with boot-strapped residuals	18.40	60.60	31.80	54.30

4.3.2 Changes in the foreign ownership restrictions

As pointed out earlier, restrictions on foreign ownership were abolished in Finland starting in 1993. This could potentially have had a huge effect on the overall integration of the market. To study this possibility we test the international CAPM using two subperiods: 1987–1992 and 1993–1998. We report adjusted *R*-squares for the regressions in Table 6. Unexpectedly the explanative power of the model does not seem to increases from the first subperiod to the other (e.g., the average *R*-square for the size portfolios is 0.201 for the first subperiod and 0.196 for the other). We also find a few surprising results: the explanative power of the international CAPM has increased strongly for the Metal & Electronincs industry (from 18.3 percent to 34.6 percent) and Food industry (from 9.0 percent to 20.0 percent), whereas it decreased for Housing & Construction industry (from 15.4 percent to 4.4 percent).

It is also interesting to know if there has been changes in the global market exposure i.e. beta. To study this possibility, we run a regression where we test whether portfolios have experienced some kind of level shifts in their global market risk exposure after 1993. In practice, we let the the beta to be a function of a dummy variable that gets a value of zero before 1993 and one thereafter ($\beta_{wmt} = b_0 + b_1 D_{1993-98}$). Table 6 presents the parameter estimates with a multivariate Wald test results. Level shift is found significant only for Banking & Other Financial industry and Metal & Electronics industri portfolios. In both cases the global risk exposure seems to increase after 1993. In general, no clear patters emerge though.

4.3.3 Missing sources of risk

One of the implications of the international asset pricing models is that local factors are irrelevant for asset prices because country-specific risks can be diversified away. Thus they should not affect stock prices significantly. If they do, it can be taken as a sign of segmentation in a particular stock market. Thus we test if local risk factors are found significant when added to the international CAPM model and whether they can even drive out the significance of the global risk factor. The results are given in Table 7. First we test if the local market risk is still priced in the market. Since local and global equity portfolios are correlated, we orthogonalized the local factor by regressing it on the global factor. Now the residual represents “pure” local source of market risk. The results show clearly that adding a local factor increases the explanative power of the model considerably. The local market beta is found multivariate significant using the Wald test statistic for both size and industry portfolios. However, the global market beta still remains significant.

Another relevant local source of risk could be liquidity. Namely, several studies have suggested that asset returns should be decreasing function of liquidity. From international investors point of view, the market-specific liquidity could be highly important determinant in their investment decisions. In empirical work the liquidity is often proxied by the bid-ask spread or trading volume. Here we use the change in the trading volume as a proxy for the liquidity risk.

Table 6. Subsample Tests of Unconditional International Capital Asset Pricing Model.

One-factor international capital asset pricing model for excess returns is tested using two sub-samples: 1987–92 and 1993–98. Adjusted R-squares are reported in the table. In addition, a model where beta is allowed to have a level shift after 1993 is tested. Coefficients for the pricing error, beta, and dummy variable are reported in the table from the OLS estimation where we have used excess returns and the Newey-West (1987) autocovariance and heteroskedasticity consistent covariance matrix with three lags. World stock market returns are calculated from the Morgan Stanley Capital International (MSCI) World Equity Index. Standard errors are given below the parameter estimates. Dummy variable gets a value of zero before 1993, and one thereafter. The sample size is 144 observations from January 1987 to December 1998.

Panel A: Size Portfolios	Average	Beta		Adj. R^2	
	Pricing error α	β_{wmt}	$D_{1993–98}$ x β_{wmt}	1987–92	1993–98
Largest	0.001	0.836*	0.102	0.342	0.344
	0.006	0.135	0.187		
2	–0.005	0.698*	0.166	0.262	0.308
	0.006	0.135	0.195		
3	–0.003	0.764*	–0.175	0.206	0.148
	0.007	0.195	0.253		
4	0.003	0.421*	0.222	0.121	0.208
	0.006	0.119	0.175		
5	–0.006	0.455*	0.018	0.135	0.112
	0.006	0.107	0.175		
Smallest	–0.007	0.360*	0.115	0.138	0.055
	0.006	0.127	0.222		
Wald-test	3.450	>99.999	3.402		
(p-value)	0.751	<0.001	0.757		
Panel B: Industry Portfolios					
Banking & Other Financial	–0.016	0.594*	0.545*	0.197	0.251
	0.008	0.227	0.273		
Forestry	–0.002	0.893*	–0.254	0.300	0.254
	0.005	0.169	0.205		
Trade & Transport	–0.004	0.640*	–0.090	0.223	0.151
	0.006	0.147	0.233		
Metal & Electronics	0.004	0.593*	0.537*	0.183	0.346
	0.006	0.153	0.206		
Food Industry	–0.003	0.447*	0.209	0.090	0.200
	0.006	0.138	0.274		
Housing & Construction	–0.015	0.555*	–0.036	0.154	0.044
	0.009	0.194	0.254		
Multi-Business	0.002	0.927*	–0.330	0.244	0.136
	0.008	0.176	0.265		
Wald-test	8.177	>99.999	14.648		
(p-value)	0.317	<0.001	0.041		

Results are also reported in Table 7. We find ten out of the thirteen portfolios to show significant exposure to the liquidity risk. All portfolios show consistently positive loadings on the liquidity risk factor. In most cases adding a liquidity risk factor increases the explanative power of the model by two-three percentage points when compared to the results in Table 3. Using a multivariate Wald test we find the liquidity factor multivariate significant for both the size and industry portfolios.

Overall we can conclude that finding local market and liquidity risks to be priced by the market implies that the Finnish stock market

Table 7. Additional Tests of Unconditional International APM Models.

Two asset pricing models are tested where the first model (segmented model) has global equity market and orthogonalized local market factors and the second model (liquidity risk model) has global equity market portfolio and a measure of local market liquidity as risk factors. Global factor returns are measured from the MSCI global equity index. Local market return is calculated from the HEX index. The liquidity measure is the change in the value of the local market trading volume. Panels A and B report the risk factor betas and Jensen's alpha using size and industry portfolio returns, respectively. All returns are in excess of the Finnish one-month Helibor rate. Parameter estimates are from the OLS estimation where we have used a Newey-West (1987) heteroskedasticity and autocorrelation consistent estimator with three lags. Standard errors are in parenthesis below parameter values. The cross-sectional Wald test statistic is also given with *p*-value in parenthesis. The sample size is 144 monthly observations from 1987 to 1998.

Time series	Segmented Model			Liquidity Risk Model			Adj. R^2	
	α_0	β_{wmt}	β_{lmt}	α_0	β_{wmt}	β_{LIQ}	Segmented	Liquidity
Panel A: Size Portfolios								
Largest	0.004*	0.875*	0.998*	–0.000	0.856*	0.042*	0.929	0.358
	0.002	0.033	0.039	0.006	0.101	0.014		
2	–0.003	0.766*	0.844*	–0.006	0.751*	0.032*	0.769	0.297
	0.003	0.049	0.066	0.006	0.109	0.013		
3	–0.001	0.683*	0.841*	–0.004	0.663*	0.041*	0.598	0.219
	0.004	0.076	0.111	0.006	0.116	0.018		
4	–0.006	0.515*	0.681*	–0.003	0.503*	0.028*	0.492	0.149
	0.005	0.082	0.110	0.006	0.092	0.010		
5	–0.005	0.459*	0.637*	–0.007	0.445*	0.029*	0.480	0.152
	0.004	0.068	0.074	0.006	0.079	0.012		
Smallest	–0.005	0.407*	0.676*	–0.007	0.394*	0.028	0.398	0.096
	0.005	0.116	0.102	0.006	0.112	0.016		
Wald-test	6.831	>99.999	>99.999	4.325	>99.999	36.551		
(p-value)	0.337	<0.001	<0.001	0.633	<0.001	<0.001		
Panel B: Industry Portfolios								
Banking & Other Financial	–0.012*	0.826*	0.961*	–0.015	0.802*	0.049*	0.561	0.230
	0.006	0.113	0.124	0.008	0.179	0.017		
Forestry	–0.001	0.774*	0.676*	–0.003	0.767*	0.025*	0.634	0.297
	0.004	0.088	0.071	0.005	0.112	0.014		
Trade & Transport	–0.003	0.597*	0.662*	–0.005	0.584*	0.028*	0.571	0.213
	0.004	0.072	0.078	0.006	0.108	0.010		
Metal & Electronics	0.008*	0.822*	0.969*	0.005	0.801*	0.044*	0.702	0.292
	0.004	0.072	0.080	0.006	0.143	0.020		
Food Industry	–0.001	0.536*	0.488*	–0.003	0.533*	0.009	0.243	0.131
	0.005	0.111	0.090	0.006	0.128	0.012		
Housing & Construction	–0.013	0.535*	0.864*	–0.011	0.523*	0.028	0.402	0.106
	0.007	0.100	0.119	0.009	0.136	0.017		
Multi-Business	0.004	0.777*	1.091*	–0.000	0.760*	0.038*	0.815	0.203
	0.005	0.100	0.083	0.008	0.126	0.019		
Wald-test	13.686	>99.999	>99.999	8.803	>99.999	31.887		
(p-value)	0.057	<0.001	<0.001	0.267	<0.001	<0.001		

has been at least partly segmented during the sample period and that a partly segmented asset pricing model could be a better characterization for the Finnish market.

5. Conclusions

This paper has investigated unconditional international asset pricing models using excess and real returns on Finnish size and industry stock portfolios. Using traditional mean-variance efficiency tests we find partial support for the international CAPM. We cannot reject the efficiency of the global equity market portfolio using size portfolios as our test assets. However, industry portfolios while showing larger variation in their exposure to global risks, reject the model. The results also show support for the international multifactor asset pricing models where we have global interest rate and Fama-French type global value premium risk factors added to the model although the explanative power of the models still remains relatively low.

This paper has concentrated on unconditional models, but there is growing evidence that the risk-premium as well as the risk sensitivities are time-varying. Therefore, a natural extension to this study is to investigate international pricing models in a conditional framework. In addition, we found strong evidence for the relevance of the local market and liquidity factors thus suggesting that a partly segmented asset pricing model could be more appropriate for the pricing of the Finnish stocks. Finally, we have also assumed that the risk of currency exchange rate changes is not relevant for the investors. A natural extension to this study would be to study currency risk in the Finnish market.

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