Finnish Economic Papers – Volume 20 – Number 1 – Spring 2007

GLOBAL MARKET AND CURRENCY RISK IN FINNISH STOCK MARKET*

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Using conditional international asset pricing models, this paper investigates whether global market and currency risks are priced in the Finnish stock market. We take the view of Finnish investors and study the pricing of the market portfolio and industry portfolios using monthly data from 1987 to 2000. The results give strong support for the pricing of the global market risk yet the local market risk is also priced suggesting mild segmentation. We find strong support for the pricing of currency risk. This suggest that Finnish investors should be concerned with the currency risk even when investing in domestic stocks. (JEL: F30, G12, G15)

1. Introduction

International aspects of asset pricing as well as many small and/or emerging stock markets have recently received increasing attention, yet many aspects still need further investigation. First, prior studies on international asset pricing models use data mainly from large markets closely integrated with global financial markets.[1] However, many small developed countries have only recently experienced full liberalization of their capital markets and many emerging countries are still in the midst of the liberalization process. Second, in many small and open economies, currency risk can play a very important role and their exchange rate mechanism often differs, e.g., from that of the USA. Third, earlier evidence suggests that predictability and autocorrelation in stock returns and the influence of local information is more evident in small equity markets. Fourth, earlier studies have usually tested international asset pricing models using only aggregated country level data.[2] However, using aggregated country indices loses many of the firm/industry cross-section differences.

This paper explores whether the global and local market risks are priced in the Finnish stock market using market and disaggregate (industry portfolio) return data from 1987 to 2000. This study extends the unconditional analysis in Vaihekoski (2000). He finds global and local market risks to be priced using unconditional beta pricing models and monthly data from 1987 to 1998. Here we use conditional asset

* *I would like to thank anonymous referees, Björn Hansson, Robert Korajczyk, Eva Liljeblom, Anders Löflund, Otto Toivanen, and seminar participants at the GSF's Joint Finance Research Seminar for their helpful comments.*

[1] *See, e.g., a survey by Stulz (1995).*

[2] *Exceptions include, e.g., Choi, Hiraki, and Takezawa (1998), Vassalou (2000), and Doukas, Hall, and Lang (1999).*

pricing models and price of risk approach that allows us to test a model also for the equity market risk premium. In addition, we study whether Finnish investors should be concerned with the currency risk even when investing in stocks from their own country.

Our primary goal is to find out whether both global and local market risks are priced in the Finnish stock market. If markets are globally integrated, investors can diversify their holdings across countries and thus the local market risk should not be systematically priced. However, like many other small yet developed stock markets, the Finnish stock market experienced a liberalization process rather recently, starting in the early 1980s. The final step in this process was not taken until beginning of 1993, when all restrictions on foreign ownership were abolished (for details, see e.g. Vaihekoski, 1997). Thus, not surprisingly, Harvey (1995a) finds the Finnish stock market to be one of the least correlated developed markets with the world equity market using data from 1988 to 1992. Nummelin and Vaihekoski (2002), on the other hand, found both the local and global market risks to be priced in Finnish asset returns using data from 1987 to 1996.

Our second goal is to study whether the currency risk is priced in the Finnish stock market. The Finnish stock market is an interesting case for this kind of studies since Finland has experienced several different exchange regimes ranging from target zones to floating exchange rate and finally membership in the EMU. In addition, the Finnish currency also experienced many speculative attacks especially in the early 1990s and the Bank of Finland was forced to defend the currency against devaluations using extremely high shortterm interest rates. Ultimately, Finnish Markka was forced to float in 1992. Later in 1995 Finland joined the European Union and in 1996, the float came to an end as Markka was inserted into the Exchange Rate Mechanism (ERM) although it still allowed for some fluctuation. Beginning in 1999, the Markka has been tied to Euro with a fixed exchange rate.

There are several reasons to believe that the currency risk is priced in the Finnish market. First, the Finnish economy has traditionally been very dependent on foreign trade.[3] On the other hand, industries differ in their export (and import) dependency suggesting different exposures to currency risk across industries. Second, the economical and political integration process within the EU and its monetary system has made Finland more closely tied to the European and global economy. Third, the Finnish government resorted many times to the devaluation of the Finnish currency during the sample period as a tool to maintain the competitiveness of Finnish industries. In fact many key Finnish industries, especially forest and metal industries, were actively promoting devaluations during the late 1980s and early 1990s. These reasons suggest that *a priori* currency risk should be priced in the Finnish stock market. Indeed, Roll (1992) as well as Antell and Vaihekoski (2007) find support for the pricing of the bilateral USD/FIM currency risk in the Finnish stock market from US investors point of view.

Finally, we study if the rewards for market and currency risks are timevarying. Small stock markets often exhibit a great deal of predictability in stock returns (Harvey, 1995b). In the Finnish stock market this has been reported by Vaihekoski (1998). From the point of view of an asset pricing model, once the predictability in the model's parameters have been accounted for, the pricing error should no longer be predictable.

The remainder of this paper is organized as follows. In the next section, the international asset pricing models of interest are presented together with their testable implications. We also consider some of the methodological and econometric questions at hand. Section 3 presents the data in this study. Section 4 presents the main empirical findings from the tests of the asset pricing models. Finally, section 5 concludes and provides our interpretation of the results together with some suggestions for further research.

[3] *Imports and exports accounted e.g. in 1997 for approximately 30 and 40 per cent of the total value of the Finnish GDP, respectively.*

2. Global market and currency risk

2.1 Theoretical background

If there are no restrictions on capital movements so that domestic investors are free to diversify internationally and foreign investors are allowed to invest in local markets, markets are said to be integrated. In this case e.g. Adler and Dumas (1983) have shown that the global value-weighted market portfolio is the relevant risk factor to consider. Assuming e.g. that investors do not hedge against exchange risks and a riskless asset exists, the conditional version of the world CAPM implies the following restriction for the nominal[4] excess returns

$$(1)\quad E\,[r_{it}|\Omega_{t-1}] = \beta_{it} E\,[r_{mt}|\Omega_{t-1}]$$

where the global market beta is

$$(2)\quad \beta_{it} = \frac{Cov(r_{it},\, r_{mt}|\Omega_{t-1})}{V\,ar(r_{mt}|\Omega_{t-1})}$$

and $E[r_{it}|\Omega_{t-1}]$ and $E[r_{mt}|\Omega_{t-1}]$ are expected returns on asset i and the global market portfolio conditional on investors' information set Ω_{t-1} available at time $t-1$. Both returns are in excess of the local riskless rate of return r_{ft-1} for the period of time from $t-1$ to t. The global market portfolio comprises all securities in the world in proportion to their capitalization relative to world wealth (see Stulz, 1995). All returns are measured in one numeraire currency.

Here we rewrite the equation (1) using the fact that the ratio $E\,[r_{mt}|\Omega_{t-1}]\; V\,ar(r_{mt}|\Omega_{t-1})^{-1}$ can be considered as the price of global market risk λ_{mt}.[5] It measures the compensation the representative investor must receive for a unit increase in the variance of the market return. Now the model gives the following restriction for the expected excess returns for assets

$$(3)\quad E\,[r_{it}|\Omega_{t-1}] = \lambda_{mt} Cov(r_{it},\, r_{mt}|\Omega_{t-1}).$$

Since the market portfolio is also a tradable asset, the model gives the following restriction for the global market portfolio's expected excess returns

$$(4)\quad E\,[r_{mt}|\Omega_{t-1}] = \lambda_{mt} V\,ar(r_{mt}|\Omega_{t-1}).$$

However, any investment in a foreign asset is always a combination of an investment in the performance of the asset and in the movement of the foreign currency relative to the domestic currency. Adler and Dumas (1983) show that if the purchasing power parity (PPP) does not hold, investors want to hedge against changes in the PPP deviations between all investable C countries with different currencies. Specifically, the risk induced by the PPP deviations is measured as the exposure to both the inflation risk and the currency risk associated with the currencies.

Assuming that domestic inflation is non-stochastic over short-periods of time, the PPP risk contains only the relative change f_{ct} in the exchange rate between the numeraire currency and the currency of country c (see, e.g., De Santis and Gérard, 1998). In this case the conditional multifactor world asset pricing model implies the following restriction for asset i's expected return in the numeraire currency

$$(5)\quad E_{t-1}[r_{it}] = \lambda_{mt} Cov_{t-1}(r_{it},\, r_{mt}) + \sum_{c=1}^{C} \lambda_{ct} Cov_{t-1}(r_{it},\, f_{ct}).$$

and similarly for the global market risk premium

$$(6)\quad E_{t-1}[r_{mt}] = \lambda_{mt} V\,ar_{t-1}(r_{mt}) + \sum_{c=1}^{C} \lambda_{ct} Cov_{t-1}(r_{mt},\, f_{ct}).$$

where λ_{mt} and λ_{ct} ct are the conditional price of market and exchange rate risks, and $V\,ar_{t-1}(.)$ and $Cov_{t-1}(.)$ are short-hand notations for conditional variance and covariance operators, all conditional on information Ω_{t-1}.

Unfortunately, the model above is intractable in practice if the number of currencies is large. Thus, one can either focus on a subset of currencies or use a more parsimonious measure for the currency risk. Ferson and Harvey (1993) show how one can use a single aggregate ex-

[4] Originally, the restriction implied by the ICAPM holds for the real excess returns, but since we are testing the model within one country, the real returns can be replaced with nominal returns (see Stulz, 1995).

[5] The price of risk is sometimes also called as reward-to-risk, compensation for covariance risk, or aggregate relative risk aversion measure.

change risk factor to proxy for the deviations from the PPP to the model. In this case, the model above collapses into a two-factor model.

If the market is partially (mildly) segmented from the global markets, the local market risk is also priced. If this is the case, the model by Errunza and Losq (1985) implies that the two-factor model has to be augmented to

$$(7)\quad E_{t-1}[r_{it}] = \lambda^g_{mt}Cov_{t-1}(r_{it}, r^g_{mt}) + \lambda_{xt}Cov_{t-1}(r_{it}, r_{xt}) + \lambda^l_{mt}Cov_{t-1}(r_{it}, r^l_{mt}).$$

where λ_{xt} is the conditional price of aggregate exchange rate risk, and r_{xt} is the return on aggregated basket of currencies. Superscripts g and l have been added to the model to emphasize the difference between global and local prices of risk and market portfolios.

2.2 *Modeling conditional expectations*

If we want to study the implications of the conditional asset pricing models in a conditional framework, we need to decide how we model investors' conditional expectations. Moreover, we have to decide what we want to model. Here, we model expected risk factor returns, and prices of risk (i.e., the ratio of risk premium to variance). Commonly used approaches to model these are the instrumental (conditioning) variables method or the time-series forecasting approach. Here we use the former method.

The instrumental variables approach presents, however, a problem because the full information set $\mathbf{\Omega}_t$ is usually not observable. Hence, we are forced to take conditional expectations on a coarser set of information $\mathbf{Z}_t \subset \mathbf{\Omega}_t$ available to us. Assuming that the information set $\mathbf{Z}_t$ fully captures the movements in the pricing model parameters or correspondingly that the full information set is generated and governed by the smaller subset of information, we can write the conditional asset pricing model by simply replacing the full information set $\mathbf{\Omega}_t$ with the subset $\mathbf{Z}_t$ (cf., Dumas and Solnik, 1995).

In addition, we have to specify a method that investors use to form their expectations from the information variables. Since in practice this is unknown to us, we have to use different econometric approaches to approximate the expectation operator (regression curve). Most of the previous research has implicitly assumed that these variables are drawn from a spherically invariant distribution, and as a result the expectation can be written as a linear regression on variables (see, e.g., Ferson, 1989).[6]

Hence we model the conditional expectation for the premium of risk factor j measured in the numeraire currency as a linear function of the local and global variables as follows

$$(8)\quad E\,[r_{jt}|\mathbf{Z}^{g,l}_{t-1}] = \mathbf{Z}^{g,l}_{t-1}\phi$$

where $\mathbf{Z}^{g,l}_{t-1}$ is a $(1 \times L)$ vector of instrumental variables' observations, and ϕ is $(L \times 1)$ vectors of linear regression coefficients, and L is the number of instrumental variables. Instrumental variables do not need to be the same for all risk factors, but typically they are the same to keep the system estimable.

Similarly, we have to decide what variables are used to model the timevariation in the prices of risks. The price of global market risk should be the same for all countries regardless of the currency used as the numeraire currency (c.f., De Santis and Gérard, 1998). Here we write it as a linear function of the global information variables

$$(9)\quad E\,[\lambda^g_{mt}|\mathbf{Z}^g_{t-1}] = \mathbf{Z}^g_{t-1}\phi$$

where $\mathbf{Z}^g_{t-1}$ is the global information set $\mathbf{Z}^g_{t-1} \subset \mathbf{Z}^{g,l}_{t-1}$ including L_g global information variables, and ϕ is a $(1 \times L_g)$ vector of coefficients. Prices of currency and local market risks are functions of global and local information variables since they are specific to the Finnish market.

2.3 *Testable implications and testing procedure*

Using the parameterization for the conditional expectations we can now test conditional pric-

[6] *Note that here it is implicitly assumed that i) the information set stays the same through time, and ii) the relation between information and expectations stays the same (i.e., coefficients are time-invariant). This, of course, is not necessarily the case.*

ing models using realized asset returns. Following previous research we utilize the generalized method of moments (GMM) approach in our analysis. The GMM estimator is efficient among the class of instrumental estimators defined by the orthogonality conditions (Greene, 1997).[7] To improve the small-sample properties of the GMM estimator, we use the iterated GMM, where the weighting matrix is found iteratively.

We begin our tests using the constant prices of risk approach. To test the world asset pricing models with K risk factors we write the following error terms

(10a) $\mathbf{u}_{1t} = \mathbf{f}_t - \mathbf{Z}^{g,l}_{t-1}\phi$

(10b) $\mathbf{u}_{2t} = \mathbf{f}_t - \lambda(\mathbf{u}'_{1t}\mathbf{u}_{1t})$

(10c) $u_{it} = r_{it} - \lambda(\mathbf{u}'_{1t} r_{it})$

where $\mathbf{f}_t$ is a $(1{\times}K)$ vector of excess risk factor returns, ϕ is a $(L{\times}K)$ coefficient matrix, r_{it} represents the realized excess return on assets i, $\mathbf{Z}^{g,l}_{t-1}$ is a $(1 \times L)$ vector of instrumental variables for conditional expected risk premium, and λ is a $(1 \times K)$ vector of the common price of risk measures.

Equation (10a) has been used to define forecast error terms $\mathbf{u}_{1t}$. These error terms have been used to define conditional variances and covariances in equations (10b) and (10c).[8] Note that in the estimation we restrict the global market risk premium to be a function of only its own variance and covariance with the currency risk factor, not with the local variance.

The asset pricing model implies that the mean error term should be conditionally zero i.e. $E[\mathbf{u}_{1t}, \mathbf{u}_{2t}, u_{it}|\mathbf{Z}^{g,l}_{t-1}] = E[\mathbf{u}_t|\mathbf{Z}^{g,l}_{t-1}] = 0$. Typically, the implementation of the error term restriction in the GMM tests uses only the linear orthogonality restriction: $\mathrm{E}[\mathbf{u}t \otimes \mathbf{Z}^{g,l}_{t-1}] = 0$, which is implied by the former condition but which is not the only implication.

The full model has $K \times L + K$ parameters to estimate and $K \times L + K \times L + N \times L$ orthogonality conditions. The model is thus overidentified in all meaningful cases and can be tested using Hansen's (1982) *J*-test statistic (test of overidentified restrictions). If the *J*-test do not reject the model outright, we can test other implications of the model. Namely, one of the implications of the pricing models is that the average pricing error should be zero. If we add an asset-specific intercept, α_i, to the model, it should be zero for all assets if the model is correct. Alpha can also be seen as a measure of mild market segmentation or as an average measure of the effects caused by other factors that are not included in the model (cf., De Santis and Gérard, 1998).

Previous research has suggested that the price of risk is time-varying. To study this case we can replace equation (10b) and (10c) with the following ones

(11a) $\mathbf{u}_{2t} = \mathbf{f}_t - (\mathbf{Z}_{t-1}\theta)(\mathbf{u}'_{1t}\mathbf{u}_{1t})$

(11b) $u_{it} = r_{it} - (\mathbf{Z}_{t-1}\theta)(\mathbf{u}'_{1t}r_{it})$

where $\mathbf{Z}_{t-1}$ is either the full or the global subset of the used information variables depending on the risk factor, θ is a $(L_g{+}L{\times}(K{-}1))$ vector of coefficients to be estimated. Now we have more parameters to estimate, but the model is still typically overidentified. The hypothesis of constant price of risk is nested in the specification as the information set includes a constant and it can be easily tested. Note that the conditional one-factor world CAPM model is also nested in the model. The specification above has, however, several shortcomings. Using linear regression (9) to model the price of risk parameters has the problem that it does not guarantee the positivity of the market price of risk. However, while this is a shortcoming of the specification, it can help to test the validity of the model, namely to help detect deviations from the positivity of the risk premium.

[7] *The GMM has several advantages that have made it popular in financial model analysis. Here, we benefit especially from the fact that the GMM does not rely upon the assumption of normally distributed asset returns; the disturbance term can be both serially dependent and conditionally heteroscedastic (MacKinlay and Richardson, 1991).*

[8] *Here we have used the fact that realized return is a sum of the expected value plus an error term. The variance term gives* $Var(r_{mt}|Z^{g,l}_{t-1}) = Var(E(r_{mt}|Z^{g,l}_{t-1}) - u_{mt}|Z^{g,l}_{t-1})$ *which is equal to* $Var(u_{mt}|Z^{g,l}_{t-1})$. *Now, the variance can be written as* $E(u^2_{mt}|Z^{g,l}_{t-1})$ *using the properties of the variance operator. Similarly, the covariance term is* $Cov_{t-1}(E[f_{jt}|\mathbf{Z}_{t-1}] + u_{jt}, r_{it}) = Cov_{t-1}(u_{jt}, r_{it})$, *since the expectation term is constant as it is taken with respect to the same information as the covariance.*

3. Data and summary statistics

The estimation period covers 168 months of data from January 1987 to December 2000. The beginning of the sample coincides with the birth of the money market in Finland and thus we have competitively determined local short-term interest rates for the whole time period. We employ continuously compounded asset returns throughout the paper since our main focus is on the temporal behavior of asset returns. All stock returns are in excess of the one-month Finnish risk free rate of return where the riskless rate of return is calculated from the one-month interbank money market rate – on the last stock market trading day of the previous month $t-1$.[9] Monthly values for all other variables are taken to be the last available daily value of the month matching the last stock market trading day in Finland.[10]

3.1 Risk factors

We employ two types of risk factors in our international asset pricing model to represent economic risks. Our first risk factor is the global market portfolio. Global market portfolio returns are proxied by the total return on the Morgan Stanley Capital International (MSCI) world equity market index, where gross dividends are invested back into the market. The Finnish stock market is not excluded from the global market index, which could be a source of multicollinearity or it could affect the expected return on the global market portfolio (see Chan, Karolyi, and Stulz, 1992). However, Finland represents less than one percent of the total global market capitalization so the problem is small and thus we follow most of the earlier studies and use the unadjusted MSCI index.[11]

The MSCI index is originally dollar-denominated. Since we take the perspective of a Finnish investor, we have to express all returns in the Finnish currency. Thus we convert the dollar-denominated return into Finnish currency using the end-of-the-month exchange rate between Finland and the USA.[12] As a result, we have the return on the world portfolio from Finnish investors' point of view. In effect, the return on the world market portfolio comprises of the value-weighted average of the individual countries' total return in Finnish currency. This total return is in turn the sum of local-currency denominated market returns and the local currency exchange rate changes against the Finnish currency.[13] As a result, the world market return is the return that Finnish investors would consider relevant in their own currency without any hedging transactions.

Our second source of risk is related to exchange rate changes. In an ideal setting we could use several currencies in our estimation. However, in practice we have to use an aggregated currency index to make the GMM estimation feasible. Here we consider the continuously compounded change in the Finnish (official) currency index calculated by the Bank of Finland.[14] The index is calculated as the trade-weighted average of freely traded currencies where weights are updated periodically (historically 3–4 times a year) to reflect changes in foreign trade.[15] A similar trade-weighted proxy has been used e.g. by Ferson and Harvey (1993).

The trade-weighted index is expected to reflect the overall, aggregate currency risk in the market. Realized returns also measure the devaluation risk in the currency market. Figure 1 shows how the currency index has developed during the sample period. The floatation decision in September 8th, 1992 can be clearly seen

[9] During 1987–1998, the Helsinki Interbank Offered Rate (Helibor) is used. After 1999, the Euribor interest rates are used.

[10] In some cases, the Finnish equity market is closed even though, e.g., the currency market is open. In those cases, the dates for the equity data is used to decide what values are picked for the other variables.

[11] In addition, the fact that the index is not corrected for cross-country holdings could be a cause for concern, but it will probably have only a negligible effect on our results (see e.g., Bansal, Hsiesh and Viswanathan, 1993).

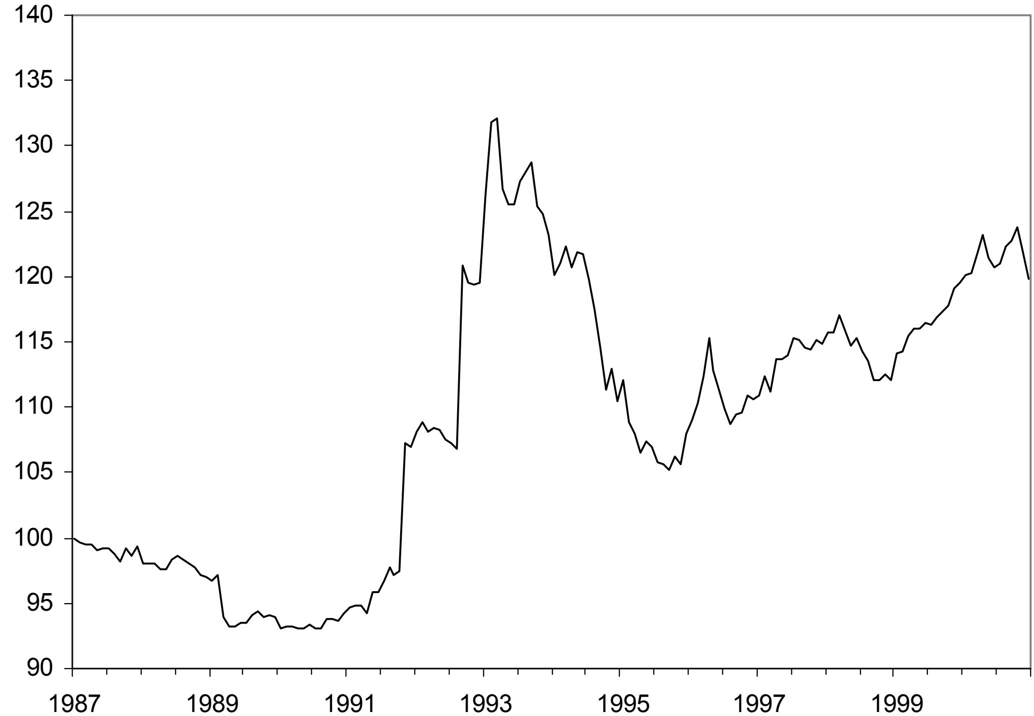
[12] During 1987–1997 we use the FIM/USD exchange rate and 1998–2000 the Euro/USD exchange rate.

[13] Ostdiek (1998) provides a more detailed discussion of the topic.

[14] Starting from 1999, after the Markka was fixed to the Euro, the currency index was renamed as Finland's competitiveness indicator.

[15] Notice that after 1999 approximately one third of the index was fixed due to the adaptation of the Euro. The biggest freely moving currencies left in the index were Swedish krona and US dollar.

Figure 1. Trade-weighted currency index for the official currency of Finland (FIM prior to 1998, Euro afterwards).



from the Figure. We can also see how the Markka started to appreciate after 1993 when the economic conditions started to improve and it became more obvious that Finland would join the EU.

Table 1 shows descriptive statistics for the excess return series. Mean and standard deviation are annualized by multiplying them with 12 and the square root of 12, respectively. The average annual nominal excess return for the world equity market portfolio from Finnish investors' point of view has been only 5.1 per cent. This is because the local risk-free interest rate was high especially during the early part of the sample period (average value over the whole sample is as high as 7.8 per cent). The currency index shows slight depreciation.

The world market returns and currency index returns show signs of nonnormality. The null of a normal distribution is rejected using the Jarque-Bera test (p-value less than 0.1 per cent). There is also evidence of significant firstorder autocorrelation in the market returns. On the other hand, the Q(12) test statistic shows evidence of long autocorrelation in four out of five currency indices.

3.2 Test assets

Initially, we test the model using one single test asset, namely the Finnish market portfolio. We do this in order to compare results with the portfolio results and with the earlier studies using a single index to represent a country e.g. by Roll (1992) and Harvey (1995b). Market portfolio returns are proxied with the logarithmic return on the value-weighted HEX yield index. The HEX index is calculated by the stock exchange and it covers all stocks quoted on the Main List of the Exchange from 1991 forward. Prior to 1991, we use the WI-index which is calculated by the Swedish School of Economics and Business Administration. For comparison, we have also included the MSCIFinland index in Ta-

Table 1. Summary statistics of monthly excess returns.

Panel A shows summary statistics for monthly returns on the value-weighted world equity index (MSCI World Equity Index), a trade-weighted currency index (Bank of Finland's official currency index), two local Finnish market indices (Helsinki Stock Exchange HEX index and MSCI Finland index), and seven industry portfolios. All returns are measured in Finnish currency in excess of the one-month Finnish risk-free rate (Helibor 1987–1998, Euribor 1999–2000). Portfolio's return is the value-weighted average of the returns for the stocks in the portfolio. Weights are updated monthly using the month-end market values from the previous month. Mean and standard deviation are annualized. Panel B gives correlation coefficients between assets' returns. The sample size is 168 monthly observations from January 1987 to December 2000.

Time series	Mean	Standard deviation	Skewness	Excess kurtosis	Normality (p-value)	Autocorrelation[a] ρ_1	ρ_2	ρ_3	Q(12)[b] (p-value)
Panel A: Summary statistics									
Risk factors									
MSCI World equity market index	0·051	0·176	−0·855	2·380	< 0·001	0·193∗	0·024	−0·072	0·194
Trade-weighted currency index	−0·065	0·058	2·811	16·634	< 0·001	0·035	0·035	−0·049	0·017∗
Local market portfolio									
HEX Finnish market index	0·121	0·266	0·029	0·474	0·450	0·265∗	0·113	0·111	< 0·001∗
MSCI Finnish market index	0·152	0·293	0·058	0·395	0·552	0·228∗	0·122	0·097	< 0·001∗
Industry portfolios									
Banking and other financial	−0·050	0·319	0·502	2·571	< 0·001	0·176∗	0·051	0·133	< 0·001∗
Forest industry	0·028	0·241	−0·094	0·517	0·346	0·124	0·054	−0·138	0·005∗
Trade and transport	−0·006	0·231	0·023	0·656	0·220	0·251∗	0·013	0·196	< 0·001∗
Metal and electronics	0·182	0·309	0·364	0·119	0·150	0·187∗	0·071	0·173	< 0·001∗
Food industry	0·022	0·292	0·174	1·823	< 0·001	0·115	0·019	0·001	0·3123
Housing and construction	−0·080	0·295	−0·199	2·035	< 0·001	0·198∗	−0·019	0·102	0·009∗
Multi-business	0·068	0·315	−0·110	0·624	0·216	0·276∗	0·049	0·041	< 0·001∗

[a] The standard error for autocorrelation coefficients with lag q is given by $\sqrt{(1 + 2\rho_1^2 + +\rho_q^2)}/\sqrt{T}$. Asterisk symbol indicates statistical significance at the 5 per cent level.

[b] $Q(12)$ is the Ljung and Box (1978) test statistic for the null that autocorrelation coefficients up to 12 lags are zero and is distributed chi-squared with 12 degrees of freedom. Asterisk symbol indicates statistical significance at the 5 per cent level.

Panel B: Correlation matrix											
Global equity market	1·000										
Trade-weighted currency index	0·381	1·000									
HEX Finnish market index	0·573	0·082	1·000								
MSCI Finnish market index	0·576	0·073	0·975	1·000							
Banking and other financial	0·423	−0·048	0·683	0·659	1·000						
Forest industry	0·511	0·172	0·700	0·669	0·483	1·000					
Trade and transport	0·396	0·021	0·661	0·592	0·571	0·593	1·000				
Metal and electronics	0·525	0·098	0·896	0·895	0·534	0·591	0·551	1·000			
Food industry	0·312	−0·006	0·467	0·413	0·410	0·397	0·497	0·327	1·000		
Housing and construction	0·313	−0·060	0·598	0·514	0·528	0·503	0·640	0·507	0·471	1·000	
Multi-business	0·403	0·022	0·752	0·734	0·522	0·620	0·590	0·591	0·449	0·523	1·000

ble 1.[16] The MSCI-Finland index includes gross dividends and is converted into FIM returns similar to the MSCI global index. It targets approximately 60 percent of the total market capitalization emphasizing liquidity and investability. MSCI indices are used frequently in international studies.

Using the aggregate market index as a test asset, one naturally looses some of the cross-sectional variation in the stocks. Therefore, we also use seven industry portfolios to test the model. The portfolios are constructed from the stocks listed on the Main List of HEX during January 1987 – December 2000 following the guidelines in Vaihekoski (2004) with some improvements.[17]

The summary statistics for the local market portfolio and the industry portfolio returns are also given in Table 1. The HEX Finnish market index shows an average of 12.1 per cent annual excess return with an annual standard deviation

[16] *MSCI Finland index is available from 1988. WI-index is used to augment the series during 1987.*

[17] *E.g., in the capitalization values of the companies. In addition, errors found in the original data series and computer program have been corrected.*

of 26.6 per cent. The MSCI-Finland index shows almost similar results. Since their correlation is 0.975, we will use the HEX index as a proxy for the Finnish stock market portfolio for the rest of the paper. Average realized market return in Finland is clearly higher than in the world market during the sample period. On the other hand, the volatility has also been much higher.

The HEX index shows a surprisingly high first order autocorrelation coefficient (0.241). It is clearly higher than that of the largest markets (typically less than 0.1). This could be driven by infrequent trading, as the coefficient is lower for the MSCI index, which includes only the largest companies. However, both series show surprisingly high persistence in autocorrelation (or predictability) of the returns on the basis of past market returns as shown by the significant $Q(12)$ test statistic.

Industry portfolios show diverse average realized excess returns as required by the tests of asset pricing models. Three out of seven industry portfolios have a negative realized average excess return during the sample period. This is probably due to the recession and banking sector crisis in the Finnish economy in the early 1990s. It is also clearly visible that the recession hit industry sectors predominantly operating in domestic markets more severely than export oriented sectors. For example, two domestic industries – banking & other financial services and housing & construction – have average realized returns of –5.0 and –8.0 per cent per annum, respectively.

Test asset returns show also signs of nonnormality. The Jarque-Bera test rejects the null hypothesis of a normal distribution for three of the assets. Furthermore, there is evidence of significant first-order serial correlation in the portfolio returns. $Q(12)$-test statistic shows evidence of higher orders of correlation.

Panel B in Table 1 presents the correlation matrix of the risk factor and test asset returns. The average correlation between the global and Finnish market portfolio returns is 0.573 during the sample period. Rolling correlation with a 24-month estimation window reveals that the correlation has varied considerably during the sample period (not reported). The correlation has increased on average, but it was surprisingly high even in the late 1980s when the Finnish stock market was partly inaccessible to foreign investors.

As expected, industry portfolios differ considerably in their correlation with the global market and currency index returns. The highest world market correlation coefficients are those for metal and electronics (0.525), and forest industries (0.511). Food industry and housing & construction exhibit the low- est correlation with the global market returns (0.312 and 0.313, respectively). The forest industry is the only portfolio showing significant correlation with the trade-weighted currency index (0.172).[18]

3.3 Information variables

In this study we employ a set of predetermined instrumental variables to track predictable time-variation in asset returns, risk exposures and the common rewards to risks. Overall, there is still no consensus on what metric to use in selecting between the variables. Simply maximizing the explanatory power of the variables raises the question of data mining. In general, the question of relevant information variables should be addressed by economic reasoning. In practice, one hopes to select theoretically justified variables that are also able to capture part of the predictability in the asset returns. Econometrically it is helpful if the specification is parsimonious.

In order to keep the model parsimonious no asset-specific information is used. The variables are chosen mostly following earlier studies. The global information set, Z^g, used to model the global market price of risk includes (1) a constant, (2) the FIM-denominated return on the MSCI world index, $R^g_{m,t-1}$, (3) a measure of the short-term yield spread (three-month Eurodollar rate minus 30-day rate), $E\$\ TS_{t-1}$, (4) a measure of world inflation (annual change in world CPI), CPI^g_{t-2}, (5) a measure of world default spread

[18] *The sample standard error for the cross-correlation coefficient can be estimated by the square root of one divided by the number of observations assuming that the population coefficient is zero and variables are normally distributed. In our sample of 168 observations, the standard error is 0.077.*

(Baa minus Aaa bond yields), $DefS^g_{t-1}$, and (6) a measure of the term structure (U.S. three-month T-bill yield minus ten-year bond yield), $USTS_{t-1}$. The world CPI measure is calculated by IMF. World default spread is calculated by Moody's for corporate bonds worldwide.[19]

The global information set is augmented with the local information set Z^l when the prices of currency and local market risks are modeled. The joint information set is also used in the statistical model for the expected risk factor returns, and conditional correlations from Finnish investors' point of view. When choosing the local information variables, one hopes that they reflect the relationship between the local and the global economy as well as the expectations related to the currency risk. Instruments that work well with one currency (e.g., the US dollar), do not necessarily work well with other currencies. Our local instrument set contains (a) a constant, (b) the return on the Finnish stock market index (HEX index), $R^l_{m,t-1}$, (c) the difference between the three-month Finnish and German money market rates (Helibor minus Fibor), $dINT_{t-1}$, and (d) a measure of the local term structure spread (12-month money market rate minus the one-month rate), $FinTS_{t-1}$.

All instrumental variables are measured with a one-period lag (world CPI measure with two lags to allow for the publication lag) so that they are in the investors' information set. In the asset pricing tests, they are also demeaned to make the interpretation of the results easier. Overall, all our instrumental variables are readily available and we have chosen this set of instruments on the basis of parsimony, and previous empirical studies (see, e.g., Ferson and Harvey, 1993; local variables, see, e.g., Vaihekoski, 1998).

Table 2 presents the summary statistics and pairwise correlation matrix of the variables. As we can see from panel B, the forecasting instruments are not especially correlated with each other. The highest pairwise correlation is 0.733 but most of the pairwise correlations are below 0.2 in absolute terms. Overall, this suggests that none of the variables are likely to be *a priori* redundant. Most of the forecasting variables appear to be nonnormally distributed. Furthermore, there is evidence of significant first and in some cases second- and third-order or even higher order serial correlation in the information variables. However, the autocorrelation coefficients appear to decay at longer lag lengths and Dickey-Fuller tests (not reported) reject non-stationarity, which suggest that the instruments are feasible for the GMM estimation.

3.4 Analysis of predictability in asset returns

To analyze if our instrumental variables are able to pick up variation in the returns of the risk factors and test assets we regress returns on global and local information variables. Adjusted R-squares from the regressions are reported in Table 3 separately for global, local, and joint instrument sets. The results show that the information variables have some predictive power over the global market portfolio (adjusted *R*-squared is 8.9 per cent) but most of the predictability is caused by the global information variables. The currency index is found predictable using both global and local information sets (adjusted *R*-squareds are 12.5 and 11.4, respectively).

Information variables are also able to predict local market and industry portfolio returns, although their explanatory power varies across portfolios. Global variables are best predictors for the metal and electronics as well as multi-business industries. This is not surprising as these industries can be considered to be the most international industries in our sample. Similarly, the local information are important especially for the domestic-oriented housing and construction industry. This result is confirmed with *F*-test statistics (*p*-values are reported in the Table) which test whether the coefficients for the information variables in question are jointly zero.

Overall, the instrumental variables seem to be able to track variation in the returns. This gives us confidence that the variables can be used to

[19] *World default spread shows correlation coefficient of 0.982 during the sample period with the more commonly used US default spread. This is not surprising since Moody's sample contained bonds mainly from the USA in the early part of period. For more information, see the documentation on Moody's website.*

Table 2. Summary statistics for information variables.

Summary statistics are given for monthly observations of the global and local information variables. The global information set includes: the world market return in FIM, the Eurodollar short-term term premium (three month Eurodollar rate minus 30-day rate), the annual percentage change in a world CPI, the world default spread in the yield difference between Moody's Baa and Aaa bonds, and the U.S. term premium (U.S. ten-year bond yield minus three-month T-bill). The local information set includes: the Finnish equity market return, the difference between Finnish and German one month money market rates (Helibor minus Fibor), and the Finnish term premimum (12-month Helibor/Euribor rate minus one-month rate). All information variables are lagged by one month (CPI measure by two months). The sample size is 168 monthly observations from December 1986 to November 2000.

Time series	Mean	Standard deviation	Skewness	Excess kurtosis	Normality (p-value)	Autocorrelation[a] ρ_1	ρ_2	ρ_3	Q(12)[b] (p-value)
Panel A: Summary statistics									
Global information variables									
World equity return	0·011	0·050	−0·876	2·589	< 0·001	0·169∗	0·013	−0·096	0·182
Eurodollar term structure	0·001	0·002	−1·236	10·323	< 0·001	0·269∗	0·181∗	0·223	< 0·001
Change in global CPI	0·138	0·072	0·432	−0·724	0·012	0·991∗	0·971∗	0·940∗	< 0·001
World default spread	0·008	0·002	0·753	−0·295	< 0·001	0·945∗	0·876∗	0·815∗	< 0·001
US term structure	−0·015	0·011	−0·225	−1·042	0·022	0·966∗	0·926∗	0·891∗	< 0·001
Local information variables									
HEX Finnish market index	0·017	0·076	0·021	0·516	0·391	0·248∗	0·091	0·088	< 0·001
Helibor-Fibor	0·024	0·027	0·808	−0·050	< 0·001	0·905∗	0·882∗	0·846	< 0·001
Finnish term structure	0·003	0·009	−1·519	4·268	< 0·001	0·653∗	0·475∗	0·322∗	< 0·001
Panel B: Correlation matrix									
World equity return	1·000								
Eurodollar term structure	−0·112	1·000							
Change in global CPI	−0·161	0·067	1·000						
World default spread	0·025	−0·181	0·277	1·000					
US term structure	0·026	−0·201	−0·492	−0·164	1·000				
HEX Finnish market index	0·562	0·029	−0·187	−0·057	−0·039	1·000			
Helibor-Fibor	−0·069	−0·152	0·418	0·733	−0·119	−0·194	1·000		
Finnish term structure	−0·185	0·263	0·069	−0·179	0·162	−0·011	−0·169	1·000	

[a] The standard error for autocorrelation coefficients with lag q is given by $\sqrt{(1+ 2\rho_1^2+ +2\rho_q^2)}/\sqrt{T}$. Asterisk symbol indicates statistical significance at the 5 per cent level.
[b] $Q(12)$ is the Ljung and Box (1978) test statistic for the null that autocorrelation coefficients up to 12 lags are zero and is distributed chi-squaredwith 12 degrees of freedom. Asterisk symbol indicates statistical significance at the 5 per cent level.

give better estimates for the expected returns as well as conditional covariances and variances than the unconditional average.

4. *Empirical results*

4.1 International asset pricing models with constant prices of risk

Even though our ultimate model is the conditional asset pricing model for partially segmented market with currency risk, we also test the world CAPM with and without the currency risk. Table 4 reports the results from the GMM estimation of the one-factor conditional world CAPM model with constant price of global market risk, i.e., equations (5) and (6). Models are estimated with an alpha parameter included for each test asset.[20] In the estimation expected returns and conditional covariances are assumed to vary through time conditional on the global and local information variables. Reported *t*-values are adjusted for heteroscedasticity and autocorrelation using Newey-West (1987) heteroskedasticity and autocorrelation consistent (HAC) covariance matrix estimator.[21]

[20] We also test a model without alpha-parameters (results not reported). The results are basically similar to those reported.

[21] The number of lags is set to three in all subsequent estimations on basis of the serial correlation patterns in the asset returns. Three is less than the recommendation that in the case of no prior, the number of lags should be close to $T^{1/4}$ which in this case with 168 observations is 3.6 (see, e.g., Hamilton, 1994). However, Andrews (1991) finds that in the neighborhood of the optimal value, variations in the lag number have little influence on the performance of the HAC estimator.

Table 3. Analysis of predictability in returns.

Excess return on risk factors, local market portfolio, and seven industry portfolios are regressed on local, global and combined information sets. The global information set includes: the world market return in FIM, the Eurodollar short-term term premium (three month Eurodollar rate minus 30-day rate), the change in a global CPI, the world default spread, and the U.S. term premium (U.S. ten-year bond yield minus three-month T-bill). The local information set includes: the Finnish equity market return, the difference between Finnish and German one month money market rates (Helibor minus Fibor), and the Finnish term premimum (12-month Helibor/Euribor rate minus one-month rate). Adjusted R-squares are reported from the regressions. F-tests are used to test the significance of the variables with p-values reported in the table. The sample size is 168 monthly observations from 1987 to 2000.

	Adj. R^2			F-test, exclude		
Return series	$Z^{g,l}$	Z^g	Z^l	$Z^{g,l}$	Z^g	Z^l
Risk factors						
Global market portfolio	0·089	0·064	−0·002	0·003	0·001	0·461
Trade-weighted currency index	0·144	0·125	0·114	< 0·001	0·016	< 0·001
Local market portfolio						
HEX Finnish market index	0·114	0·111	0·082	< 0·001	0·030	0·001
Industry portfolios						
Banking and other financial	0·086	0·094	0·028	0·004	0·011	0·636
Forest industry	0·053	0·063	0·001	0·033	0·020	0·765
Trade and transport	0·052	0·055	0·030	0·034	0·121	0·457
Metal and electronics	0·105	0·106	0·073	0·001	0·058	0·416
Food industry	0·050	0·059	0·000	0·039	0·022	0·711
Housing and construction	0·054	0·028	0·054	0·031	0·414	0·065
Multi-business	0·133	0·126	0·050	< 0·001	0·001	0·228

In panel A the only test asset has been the Finnish market portfolio (proxied by the HEX-index). Results are shown for three different models. The first model includes only the global market risk factor. The second includes also the currency risk factor. The third model adds the local market risk into the model. Notice that the risk factor returns have also been assumed to follow the model in question except in the case of the partially segmented model, where the global risk factor has been priced according to the two-factor model (global market and currency risk).

The results show that the price of global market risk is significantly different from zero. It is also found positive (14.396, 15.146, and 7.644 for the three models, respectively) as suggested by the asset pricing model. It is, however, clearly higher than what has been traditionally found acceptable by the finance theory (less than five) and what earlier empirical studies have found. For example, De Santis and Gérard (1998) found the constant price of global market risk to be 2.79. The currency risk and the local market risk are found to be significantly priced if included in the model.

The test of overidentifying restrictions reject all but the three-factor model (p-value 0.433). On the other hand, the α-parameter is in all cases significantly different from zero. This has been traditionally interpreted as a rejection for the model. Altogether the results seem to suggest that the world CAPM alone is not adequate to price the Finnish market portfolio even if the currency risk is added to the model.

Panel B reports results for the industry portfolios. A system of equations is estimated for all assets at the same time with separate alphas for each asset. The results are similar to those in panel A with a couple of differences. First, the test of overidentifying restrictions does not reject any one of the models. There can be several explanations for this. The J-statistic is more likely to accept the model if the industry portfolios have lower pricing error on average than the HEX-index or the variation in the errors is

Table 4. Conditional international APMs with constant prices of risks.

The table presents constant (unconditional) price of risk estimates and average pricing errors (alphas) for three different versions of the conditional international asset pricing model. Results are estimated using the GMM. Expected risk factor returns and covariances are conditional on the global and local information set $Z^{g,l}$. Panels A and B report results using the Finnish market portfolio and seven industry portfolios as test assets, respectively. Used risk factors are also modeled according to the tested model. Model I is the basic one-factor world CAPM. Model II is two-factor world APM where the currency risk factor has been added to the model. Model III is a segmented three-factor APM where the local market risk has been added to the model. All returns are in excess of Finnish risk-free rate. *T*-values are reported below the parameter estimates in parenthesis. They have been adjusted to autocorrelation and heteroscedasticity using Newey-White (1987) correction. Coefficients significantly (5 %) different from zero are marked with an asterisk (*). An overall test of overidentifying restrictions (so-called *J*-test) is also provided in the table. The degrees of freedom (d.o.f.) is the number of orthogonality conditions minus the number of parameters. The sample size is 168 monthly observations from 1987 to 2000.

Panel A: Finnish market portfolio as test asset

	Model I	Model II		Model III		
Price of risk:	λ_m	λ_m	λ_{fx}	λ_m	λ_{fx}	λ_l
	14·989*	15·146*	−46·161*	7·644*	−20·507*	31·207*
	(9·189)	(10·325)	(−10·798)	(4·234)	(−6·259)	(4·973)
Average pricing error:	α_m		α_m			α_m
HEX-index	−0·012*		−0·014*			0·004
	(−2·373)		(−3·670)			(0·355)
J-test:	34·771		37·212			23·467
p-value	0·004		0·041			0·433
Degrees of freedom:	16		24			23

Panel B: Industry portfolios as test assets

	Model I	Model II		Model III		
Price of risk:	λ_m	λ_m	λ_{fx}	λ_m	λ_{fx}	λ_l
	12·729*	4·780*	−17·298*	2·508*	−18·756*	0·412*
	(23·969)	(16·348)	(−34·425)	(7·414)	(−32·427)	(7·327)
Average pricing error:	α_i		α_i			α_i
Banking and other financial	−0·018*		−0·009*			−0·006*
	(−15·391)		(−17·699)			(−8·502)
Forest industry	−0·016*		−0·003*			−0·004*
	(−13·867)		(−6·621)			(−6·858)
Trade and transport	−0·014*		−0·005*			−0·004*
	(−13·867)		(−10·212)			(−8·370)
Metal and electronics	−0·020*		−0·005*			−0·005*
	(−21·487)		(−10·933)			(−9·547)
Food industry	−0·012*		−0·005*			−0·007*
	(−7·925)		(−7·460)			(−8·628)
Housing and construction	−0·013*		−0·007*			−0·007*
	(−10·824)		(−15·188)			(−11·711)
Multi-business	−0·017*		−0·006*			−0·008*
	(−12·678)		(−10·502)			(−9·971)
J-test:	51·837		54·319			53·737
p-value	0·863		0·940			0·989
Degrees of freedom:	64		72			80

smaller on average. It also implies that the model is better suited to explain the cross-sectional variation in the industry portfolio returns than the aggregated market return. Second, the pricing errors (alphas) are smaller than for the whole market, although they are still significantly different from zero.

In conclusion, the results in Table 4 suggest even though none of the models can price the test assets perfectly, the three factor model is the most appropriate model of the models analyzed here for the pricing of stocks in the Finnish equity market.

4.2 International asset pricing models with time-varying prices of risk

Earlier research has found the prices of market and currency risk to be timevarying. Therefore we re-test our three-factor asset pricing model using a specification where the prices of risk are allowed to vary through time. Price of global market risk in conditioned on the global information variables whereas prices of currency and local market risk are conditioned on global and local information variables. As before, we first estimate the model using the Finnish market portfolio and then using the industry portfolios. Results are reported in panels A and B of Table 5.

The results in both panels are surprisingly similar – again the *J*-statistic rejects the models when only one test asset is used and accepts the models when several assets are used. The reported parameter estimates are for the information variables used to model the prices of risk. Since the variables are demeaned, we can interpret the constant as the unconditional mean for the price of risk. The unconditional prices of global market and currency risk are found significantly different from zero. The mean price of global market risk is even within the region deemed acceptable in the existing literature when the model is tested using the industry portfolios.

The unconditional price of local risk is not found significant. However, *LR*-test rejects the hypotheses of constant rewards to risk indicating that all prices of risk are time-varying (reported in Table 5) and different from zero (not reported).[22] These results support our earlier finding that all of our three sources of risk are priced.

We can also use the parameter estimates and the conditioning variables to calculate the conditional risk premium for the assets as well as the prices of risk. Using the estimates from Panel A in Table 5, the three-factor model implies a monthly equity market risk premium of 0.734 per cent (8.81 % per annum) for Finland.[23] It is slightly higher than what researchers have suggested as market premium but nevertheless much more realistic forecast than the historical risk premium (12.1 % per annum). One reason for the high implied risk premium could be that we did not specify any model for the covariances or variances and their realizations may differ from investors' expectations. In addition, the model might not account properly for the true nature of currency risk as suggested by Berglund and Löflund (1996).

Figure 2 shows how the prices of risks have evolved during the sample period. Notice that we did not have any restrictions on the price of global market risk, so it can also show negative values.[24] However, the point estimates of the process are always subject to estimation error and thus, more interesting insights can be made from the general trend in the process. Price of global market risk is higher than average especially after early-1990s. Naturally, we can only speculate on the reasons. This part of the sample period suffered from slow growth in the global economy which could explain the higher price of risk. This study and De Santis and Gérard (1998) have partly overlapping sample periods (1987–1994) and the trends seem to be

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[22] LR-tests in GMM estimation is first done without the restriction after which the estimation run again with the restriction and the weighting matrix from the unconstrained model. Only one iteration over the parameters is allowed. The reported test statistic is the difference in the overall J-statistics. The degrees of freedom are equal to the number of restrictions

[23] Note that the result has to be interpreted as an average of the conditional implications, since our GMM system is actually a test of the unconditional implications of the conditional models as pointed out by De Santis and Gérard (1998).

[24] We tried to force the positivity using the exp($\mathbf{Z}\theta$)-formulation similar to Bekaert and Harvey (1995), but the estimation became too unstable.

Table 5. Conditional international APM with time-varying prices of global, currency, and local risk.

The table presents conditional price of risk estimates for the three-factor conditional international asset pricing model. Results are estimated using the GMM. The sources of risk are global and local market risk, and currency risk. Expected risk factor returns and covariances are conditional on global and local information variables. Global price of market risk is conditional on global information variables, whereas the local market and currency risks are conditional on global and local information variables. Variables are explained in the text. Panel A reports results using Finnish market portfolio as the test assets. Panel B uses seven industry portfolios as test assets. Used risk factors are also modeled according to the pricing model. All returns are in excess of local risk-free rate. *T*-values are reported below the parameter estimates in parenthesis. They have been adjusted to autocorrelation and heteroscedasticity using Newey-White (1987) correction. Coefficients significantly (5 %) different from zero are marked with an asterisk (*). An overall test of overidentifying restrictions (so-called *J*-test) is also provided in the Table. The degrees of freedom (d.o.f.) is the number of orthogonality conditions minus the number of parameters. Separate LR-tests on parameters are also reported. The sample size is 168 monthly observations from 1987 to 2000.

	Parameter estimates								
		Global variables					Local variables		
	Constant	$R^g_{m,t-1}$	$E\$ \, TS_{t-1}$	CPI^g_{t-2}	$DefS^g_{t-1}$	$USTS_{t-1}$	$R^l_{m,t-1}$	$dINT_{t-1}$	$FinTS_{t-1}$
Panel A: HEX-index as test asset									
Price of world market risk	10·930*	0·559	−53·734*	−1·132	16·851	−5·628			
	(3·614)	(0·727)	(−3·172)	(−1·735)	(0·934)	(−1·380)			
Price of local market risk	−2·156	−0·326	−1·383	4·784*	1·333*	3·814	−0·642*	10·302	−4·646*
	(−1·495)	(−1·564)	(−1·610)	(2·414)	(3·470)	(0·506)	(−2·582)	(1·271)	(−3·106)
Price of currency risk	−76·786*	−3·333*	−7·558	−30·888*	3·583	152·967*	0·363	−138·430	4·763
	(−5·125)	(−2·115)	(−1·286)	(−4·131)	(1·484)	(2·009)	(0·198)	(−1·708)	(0·425)
Test of overidentifying restrictions:	$J = 9.209 \sim \chi^2_3$, p-value $= 0.027$								
LR-test on price of global market risk is constant:	$LR = 14.930 \sim \chi^2_5$, p-value $= 0.011$								
LR-test on price of local market risk is constant:	$LR = 24.147 \sim \chi^2_8$, p-value $= 0.002$								
LR-test on price of currency risk is constant:	$LR = 18.809 \sim \chi^2_8$, p-value $= 0.016$								
Panel B: Industry portfolios as test assets									
Price of world market risk	2·754*	−0·116	−30·941*	−1·031*	−13·888*	−8·237*			
	(3·194)	(−0·770)	(−7·573)	(−7·806)	(−2·518)	(−9·070)			
Price of local market risk	0·350	−0·631*	−1·665*	5·013*	1·083*	−3·185*	−0·389*	25·098*	−2·441*
	(0·873)	(−8·171)	(−7·299)	(8·018)	(10·142)	(−2·090)	(−5·002)	(9·948)	(−5·396)
Price of currency risk	−44·576*	−2·914*	−14·214*	−22·624*	4·754*	109·126*	1·472*	164·382*	−6·957*
	(−11·935)	(−6·767)	(−8·397)	(−8·381)	(5·670)	(5·142)	(2·369)	(7·539)	(−2·510)
Test of overidentifying restrictions:	$J = 47.479 \sim \chi^2_{66}$, p-value $= 0.958$								
LR-test on price of global market risk is constant:	$LR => 99.999 \sim \chi^2_5$, p-value $= < 0.001$								
LR-test on price of local market risk is constant:	$LR => 99.999 \sim \chi^2_8$, p-value $= < 0.001$								
LR-test on price of currency risk is constant:	$LR => 99.999 \sim \chi^2_8$, p-value $= < 0.001$								

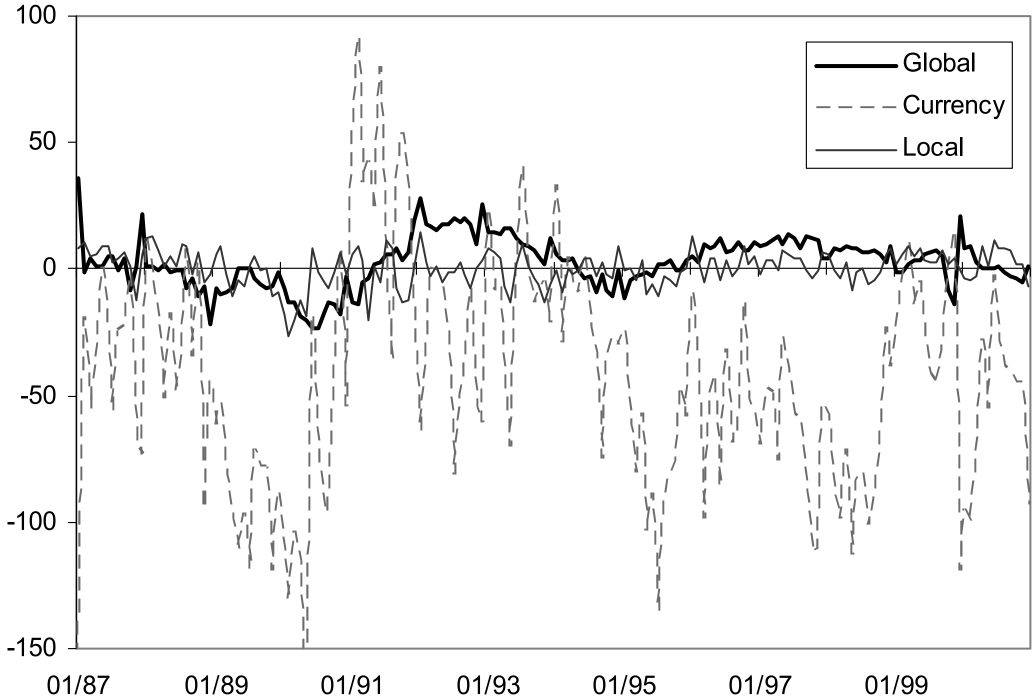
quite similar. The price of currency risk shows high time-variation as suggested by earlier research. The preceding buildup period to the floatation decision in the September 1992 can be clearly seen from the Figure.

In summary, the results show strong support for the pricing of global and local market risk as well as currency risk in the Finnish stock market. Thus, Finnish investors should take the currency risk into consideration even when investing in their own country. The results also show evidence that the rewards to taking these risks are time-varying and predictable using the selected information variables. The result are in line with those of De Santis and Gérard (1998) and others.

4.3 Additional tests

We also conduct several robustness tests. First, we check that the GMM estimation converges to the global minimum using different starting values. Second, we run the main GMM estimations using the identity matrix as the weighting matrix and with only one iteration to make sure that the model is not accepted while some of the assets are clearly not priced by the model. The results (not reported) do not change any of the

Figure 2. Time series development of the conditional prices of global and local market risk as well as currency risk (1987/1 –2000/12).



main results. Finally, we test the implication that the pricing errors should not be predictable using information variables. The predictability observed in the asset returns has disappeared from the residuals. This increases our faith in the results.

The January-effect is a well-known anomaly in the stock market. It is also evident in the Finnish stock market. If we run a regression of portfolio returns on a constant and a January dummy we get the coefficient for the January cross-sectionally significant using a Wald-test (p-value 0.013). This is a sign of inefficiency unless the asset pricing model can explain the higher return in January with higher risk. Now, we test if the three-factor model has been able to pick up this January-effect by running a regression on the residuals from the model. The results show that the model has been able to explain part of the January-effect as the January-coefficient is no longer cross-sectionally significant (p-value for a cross-sectional Wald-test statistic is 0.311) even though it is still significant for some of the portfolios.

5. Concluding comments

In this paper we study the pricing of Finnish market and industry portfolios using monthly return data from 1987 to 2000. We take the view of a Finnish investor and study whether global, local, and currency risk are priced in the market using conditional international asset pricing models. The Finnish stock market is an interesting test laboratory as it is small, volatile, and relatively recently liberalizated. Moreover, the Finnish currency has gone through various exchange regimes ranking from target zones to floating exchange rate and finally conversion to the Euro.

Similar to Vaihekoski (2000), we find strong support for the pricing of global market risk. On the other hand, the results suggest that even though investors can diversify globally, the market specific local risk is still priced and important. This implies that we cannot ignore neither global nor local sources of risk at least as far as the Finnish market is concerned. This result is in line with Nummelin and Vaihekoski (2002).

We also test whether the currency risk is priced in the Finnish stock market. The results show strong support for the exposure to currency risk in asset prices. This indicates that the global and local market risks alone are not sufficient to explain asset returns. This implies that Finnish investors should take the currency risk into account even when investing in domestic companies. Similar to De Santis and Gérard (1998) we also find strong support for the time-variation in the prices of risk and especially with respect to the currency risk.

Taken together, the results suggest that the asset pricing model for mildly segmented markets originally by Errunza and Losq (1985) could be the most appropriate one for the Finnish equity market when augmented with the currency risk. The rewards to risk (prices of risk) are time-varying and one should account for their variation in empirical models. They are also found predictable using the selected information variables offering investors a possibility to benefit from the conditional approach.

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