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ABSTRACT

We investigate the role of currency risk on stock markets in two interlinked Nordic countries exhibiting a gradual move from fixed to floating exchange rates. We apply the Ding and Engle (2001) covariance stationary specification in a multivariate GARCH-M setup to test a conditional international asset pricing model. Using a sample period from 1970 to 2009, we find that the currency risk is priced in both stock markets as well as the price to be lower after the flotation of the currencies. We also find the cross-country exchange rate shock from Finland to affect the price of currency risk in Sweden, but not vice versa. Finally, we discuss some of the potential issues in applying multivariate GARCH-M specifications in tests of asset pricing models.

JEL Classification: G12; G15

Keywords: conditional, international asset pricing model, currency risk, devaluation, multivariate GARCH-M, Finland, Sweden

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1 INTRODUCTION

During the past few decades, foreign investments have become easier and more cost efficient to conduct. The general liberalization of administrative and legal restrictions on the financial markets has provided investors a much larger investment opportunity set than ever before. As a part of this development, many developed countries have abandoned fixed exchange rate systems and moved towards market-determined floating rates and abolished foreign-exchange controls. However, there are still many emerging countries with currencies that are either fixed or tied to certain target zones.

Since the pricing of currency risk in the stock market is still a somewhat controversial issue, many papers explore the role of currency risk in asset pricing, For example, using data from large countries, De Santis and Gérard (1998) conclude that the time variation in the risk premium could explain why the unconditional models are unable to detect highly time-varying currency risk. Antell and Vaihekoski (2007) also find support for the pricing of currency risk in Finland, but they conclude that the basic time-varying price of currency risk approach does not necessarily fit countries with changing currency regimes. Especially, the devaluation risk needs more careful consideration.

In this paper we study the pricing of global and local market risks, and in particular currency risk on the Finnish and Swedish stock markets. This study extends the analysis in Antell and Vaihekoski (2007) in a number of ways. First, we add Sweden into the analysis and extend the sample period by more than four years. Both Finland and Sweden are export oriented countries known to have used competitive devaluations. This gives us a unique chance to study cross-country effects in currency risk. Second, we test for the effect of fixed and floating currency regimes on the pricing of currency risk, as both the Finnish and Swedish currencies were first pegged against a currency index within a pre-specified band but were both forced to let their currencies float almost at the same time in 1992. Finally, we discuss some of the practical caveats in using Ding and Engle (2001) GARCH specification within the framework of De Santis and Gérard (1998) as it has become popular in tests of asset pricing models (see, e.g., De Santis et al. 2003; Gérard et al. 2003; Barr and Priestley, 2004).

Overall, we believe the institutional features and the particular sample period make the Finnish and Swedish stock markets unique test laboratories for currency risk within the conditional international asset pricing framework. Including two rather similar, yet in many ways different countries allows also for interesting comparison between the countries. Our primary goal is to explore how the currency risk is priced in these stock markets. In particular, we study the role of the exchange rate mechanism. Second, we study how Finland and Sweden differ in their pricing with respect to local sources of risk. The results can shed light on the role of currency risk and local risk on the pricing of stocks in countries that are currently emerging from segmentation and also restricting the free valuation of their currencies (e.g., Eastern European new EU members, Russia, and China).

The remainder of the paper is as follows. Section 2 presents the theoretical background and research methodology. Section 3 gives a short introduction to the history of Finnish and Swedish currency policy and presents the data in this study. Section 4 shows the empirical results. Section 5 concludes and offers some suggestions for further research.

2 RESEARCH METHODOLOGY

2.1 Theoretical background

If capital markets are economically fully integrated, the expected return is driven by the same pricing model with a common set of risk factors with common risk premia in all countries. Return differences are exclusively explained by differences in the exposure to the risk factors. Suppose the correct model is given by the one-factor market model or the CAPM. Then, as shown by Adler and Dumas (1983), the expected return is driven by the exposure to the value-weighted world equity benchmark portfolio. In this case the conditional world CAPM is determined by

$$E[r_{i,t+1}|\Omega_t] = \beta_{i,t+1}(\Omega_t)\, E[r_{m,t+1}|\Omega_t], \tag{1}$$

where $E[r_{i,t+1}|\Omega_t]$ and $E[r_{m,t+1}|\Omega_t]$ are expected excess returns on asset *i* and the global market portfolio conditional on investors' information set Ω_t available at time *t*. All returns, including the risk-free rate, are measured in a common numeraire currency. Since the conditional beta is defined as $Cov(r_{i,t+1},r_{m,t+1}|\Omega_t]Var(r_{m,t+1}|\Omega_t)^{-1}$, we can use equation (1) to define the ratio

$E[r_{m,t+1}|\Omega_t]Var(r_{m,t+1}|\Omega_t)^{-1}$, i.e., the conditional price of global market risk $\lambda_{m,t+1}$.[1] It measures the compensation the representative investor must receive for a unit increase in the variance of the market return (see Merton, 1980). Now the model is given by

$$E[r_{i,t+1}|\Omega_t] = \lambda_{m,t+1} Cov(r_{i,t+1}, r_{m,t+1}|\Omega_t). \tag{2}$$

Model (2) is applicable for any asset i, and hence also for the market portfolio, in which case the model is

$$E[r_{m,t+1}|\Omega_t] = \lambda_{m,t+1} Var(r_{m,t+1}|\Omega_t), \tag{3}$$

where $Var(\cdot|\Omega_t)$ is the conditional variance of the market return. However, if some assets deviate from pricing under full integration, their risk-adjusted return will differ from the global CAPM. If this is the case, the market price of global risk should be the same for all assets everywhere, after adjusting for the costs arising from the barrier constraints. Errunza and Losq (1985) suggested including the local market portfolio as an additional source of risk in the pricing equation. Further, keeping in mind that an international investment is a combination of the direct investment into the asset itself and an indirect investment into the foreign currency, the conditional expected return for asset i can be stated as

$$\mathrm{E}_t\left[r_{i,t+1}\right] = \lambda^{w}_{m,t+1}\mathrm{Cov}_t(r_{i,t+1}, r^{w}_{m,t+1}) + \sum_{c=1}^{C}\lambda_{c,t+1}\mathrm{Cov}_t(r_{i,t+1}, f_{c,t+1}) + \lambda^{l}_{m,t+1}\mathrm{Cov}_t(r_{i,t+1}, r^{l}_{m,t+1}), \tag{4}$$

where $\lambda^{w}_{m,t+1}$, $\lambda^{l}_{m,t+1}$, and $\lambda_{c,t+1}$ are the conditional prices of world and local market risk, and exchange rate risk for currency c.[2] The conditional variance and covariance are given by $Var_t(\cdot)$ and $Cov_t(\cdot)$. Note that the price of currency risk does not need to be positive. However, including a larger set of currencies in the model might become infeasible. In this case one can focus on a subset of currencies. Alternatively, following Ferson and Harvey (1993) and Harvey (1995), one could use an aggregate currency risk factor, in which case the model would boil down to a three-factor model.

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

2 See, e.g., De Santis and Gérard (1998).

2.2 Empirical formulation

To transform the general conditional asset pricing framework into a tractable empirically testable formulation we employ a multivariate GARCH-in-Mean specification similar to De Santis and Gérard (1998) to model the conditional expectations, covariances, and variances. The starting point is Antell and Vaihekoski (2007) who study US investors investing domestically, and in one foreign country, i.e., Finland. Our framework is revised by the inclusion of Sweden as an additional stock market to invest in. Both markets are in the Nordic region, and exhibit somewhat similar, yet in many ways different characteristics.

We estimate the model using returns for four test assets: the world, U.S., Finnish, and Swedish equity market portfolios. Also the currency returns are modeled. The U.S. market is included to compare the results with De Santis and Gérard (1998), and Antell and Vaihekoski (2007). We employ the following model for excess returns in USD

$$r_{m,t+1}^{w} = \lambda_{m,t+1}^{w} h_{t+1}^{w} + e_{m,t+1}^{w}, \tag{5}$$

$$r_{m,t+1}^{US} = \lambda_{m,t+1}^{w} h_{t+1}^{US,w} + \lambda_{m,t+1}^{US} h_{t+1}^{US} + e_{m,t+1}^{US}, \tag{6}$$

$$r_{m,t+1}^{FIN} = \lambda_{m,t+1}^{w} h_{t+1}^{FIN,w} + \lambda_{m,t+1}^{FIN} h_{t+1}^{FIN} + \lambda_{t+1}^{FIM} h_{t+1}^{FIN,FIM} + e_{m,t+1}^{FIN}, \tag{7}$$

$$r_{m,t+1}^{SWE} = \lambda_{m,t+1}^{w} h_{t+1}^{SWE,w} + \lambda_{m,t+1}^{SWE} h_{t+1}^{SWE} + \lambda_{t+1}^{SEK} h_{t+1}^{SWE,SEK} + e_{m,t+1}^{SWE}, \tag{8}$$

$$r_{t+1}^{FIM} = \lambda_{m,t+1}^{w} h_{t+1}^{FIM,w} + \lambda_{t+1}^{FIM} h_{t+1}^{FIM} + e_{t+1}^{FIM}, \tag{9}$$

$$r_{t+1}^{SEK} = \lambda_{m,t+1}^{w} h_{t+1}^{SEK,w} + \lambda_{t+1}^{SEK} h_{t+1}^{SEK} + e_{t+1}^{SEK}, \tag{10}$$

$$\boldsymbol{\varepsilon}_{t+1} \sim \mathrm{IID}(0, \mathrm{H}_{t+1}),$$

where lambdas are the conditional prices of risk and $\boldsymbol{\varepsilon}_{t+1}$ is a 6×1 vector of stacked innovations, i.e., $\boldsymbol{\varepsilon}_{t+1} = [e_{m,t+1}^{w}\ e_{m,t+1}^{US}\ e_{m,t+1}^{FIN}\ e_{m,t+1}^{SWE}\ e_{t+1}^{FIM}\ e_{t+1}^{SEK}]'$. H_{t+1} is the variance-covariance matrix. Equations (5)–(10) are the empirical counterparts to the theoretical equations. Note that we have assumed that the currency risk premium is not a function of the local market risk. Further, the currency component enters only the own stock market and the currency's own equation, i.e. FIM (SEK) currency risk enters the Finnish (Swedish) stock market equation, and the own FIM (SEK) equation.

The six-variate (co)variance process of ε_{t+1} can be modeled in numerous ways. The volatility of financial assets often shows clustering, time-variation, asymmetry, and non-normality. Specifications in the family of (generalized) autoregressive conditional heteroskedasticity (GARCH) are inclined to model these stylized features. The problem with many specifications is their intractability at estimation, especially the large number of parameters to be estimated. This is especially so for the unrestricted multivariate GARCH of Bollerslev et al. (1988). Other problems include getting the variance process stationary and the variance matrix positive definite. Many of the problems are avoided by the Baba, Engle, Kraft and Kroner (BEKK) formulation set forth by Engle and Kroner (1995):

$$\mathbf{H}_{t+1} = \mathbf{C}'\mathbf{C} + \mathbf{A}'\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t'\mathbf{A} + \mathbf{B}'\mathbf{H}_t\mathbf{B}, \tag{11}$$

where C is an upper-triangular matrix, and A as well as B are 6×6 non-symmetric parameter matrices. Taken the bivariate case as an example, the matrices can be written as follows:

$$\mathbf{C} = \begin{bmatrix} c_{11} & c_{12} \\ 0 & c_{22} \end{bmatrix}, \ \mathbf{A} = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}, \text{ and } \mathbf{B} = \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix}. \tag{12}$$

Specification (11) allows for rich dynamics and a positive-definite covariance matrix. However, without restrictions on the parameter space, the number of variance parameters in a 6×6 system is still 21+36+36 = 93. To further limit the parameter space we impose diagonality restrictions on matrices A and B, and use the covariance stationary specification of Ding and Engle (2001):

$$\mathrm{H}_{t+1} = \mathrm{H}_0 \circ (\mathrm{ii}' - \mathrm{aa}' - \mathrm{bb}') + \mathrm{aa}' \circ \boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}'_t + \mathrm{bb}' \circ \mathrm{H}_t, \tag{13}$$

where $\circ$ denotes the Hadamard (or Schur) element-by-element matrix multiplication operator, and a and b contain the diagonal elements of A and B, respectively. H_0 is the unconditional variance-covariance matrix. The number of elements to estimate is now 6+6 = 12.

The estimation is conducted by quasi-maximum likelihood (QML), in which robust standard errors are computed as stated by Bollerslev and Wooldridge (1992). These can be used to calculate robust *t* statistics and Wald statistics. Provided that the conditional mean and conditional variance are correctly specified, QML yields consistent and asymptotically normally distributed parameter estimates even if the underlying distribution is non-normal. The Berndt–Hall–Hall–Hausman (BHHH) algorithm is used for the optimization.

Next, we have to decide on a model for the coefficients of price of risk. A straightforward choice is a linear representation, often used in previous research (e.g. De Santis and Gérard, 1998). For example, the linear specification of the price of world risk is

$$\lambda_{m,t+1}^{w} = Z_t^{w}\kappa_w \,, \tag{14}$$

where Z_t^w is an $1 \times L$ vector of conditioning variables (a subset of investors' information set Ω_t) and κ_w an $L \times 1$ vector of coefficients. The number of global variables is L, including the constant.[3] The models for local market risk and currency risk are modeled in a similar, linear fashion. The information sets are given by Z_t^l and Z_t^c, respectively. For the local market risk we use a combination of global and market specific variables, while currency risk is modeled by currency specific variables. Further, to study the effect of the floating decision in 1992 on the price of currency risk, we add an indicator variable DFLO for the post-floating period. Using multiplicative indicator variables for all information variables, we get the following specification for the price of currency risk for currency c:

$$\begin{aligned}\lambda_{t+1}^{c} = \kappa_0^c + \kappa_1^c DFLO^c + \kappa_2^c Z_{t,1}^c + \ldots + \kappa_{1+L_{fx}}^c Z_{t,L_{fx}}^c \\ + \kappa_{2+L_{fx}}^c DFLO^c Z_{t,1}^c + \ldots + \kappa_{1+2L_{fx}}^c DFLO^c Z_{t,L_{fx}}^c\end{aligned} \;. \tag{15}$$

The specification in (15) includes all multiplicative terms, but in the estimation some of the terms are dropped to keep the system tractable especially if there are reasons to believe that the role of a particular variable does not differ before and after the floating.

3 DATA

Our estimation period covers 474 months of data from March 1970 to August 2009. The beginning of the sample period matches that of Antell and Vaihekoski (2007) but it extends more

3 A linear representation might yield a negative price of risk. If we wish to ensure that it is always positive, one could use the exponential function for the right-hand side of equation (14).

than four years beyond, including the worldwide financial crisis that peaked in Fall 2008 and Winter 2009. We take the view of a US investor. Thus, all returns are measured in US dollars in excess of U.S. investors' risk-free return. The risk-free rate for month $t+1$ is measured as the one-month holding period return calculated from the US Treasury bills at the end of month t. The risk free rate series is taken from Ibbotson SBBI (2009) and converted to continuously compounded return.[4] We use continuously compounded asset returns throughout the paper. All returns are in percentage form.

3.1 Case: Two emerging Nordic countries and their Foreign Exchange Policy

The Finnish currency, Markka (FIM) was established in 1860 under the autonomy from Russia, while the current version of the Swedish currency, Krona (SEK), was established in 1873 (Jonung, 2000). Both currencies were tied to the gold standard at a fixed rate. Historically, both Finland and Sweden have deployed a fixed exchange rate policy, tying their currencies to gold, the USD, or some exchange rate index. However, the central banks have fairly often been forced to loosen up that policy, making devaluations (and occasionally also revaluations).

Panel A of Table 1 shows the Finnish and Swedish currency regimes since their inception to the Bretton Woods system. The FIM joined Bretton Woods in 1949, while the SEK joined 1951 which tied the currencies against the USD. For a while the Nordic currencies experienced a relatively calm period along with most of the rest of the world. However, the beginning of the 1970s changed everything as the USA unilaterally terminated the convertibility of the USD to gold in August 1971. After December 1971 FIM and SEK were determined under the Smithsonian agreement until the first half of 1973 after which both currencies were pegged to a trade-weighted currency index, first unofficially and later officially with a fixed fluctuation range. In the case of the Krona, the U.S. dollar had double weight.[5]

From 1970 to 1990 both currencies experienced several devaluations and a few occasional revaluations. See Panel B of Table 1. As a result, the value of FIM and SEK decreased during the sample period especially against the USD. In many cases, a devaluation decision in the other country sparked a similar devaluation in the other. In fact, Sweden and Finland at times accused

4 Values for 2009 are taken from Kenneth R. French's web-site.

5 For more information on the history of SEK, see Bohlin (2010).

each other for using devaluations as tools to improve their export industries' (especially metal and forestry) competitive position.

From the beginning of 1991, both FIM and SEK were linked to the European Currency Unit (ECU) with fixed rate. However, after several speculative attacks in September 1992, Finland was forced to let its currency floating. Sweden had to follow two months later in November 1992. Soon afterwards, both started to strengthen against the USD. In October 1996 FIM became part of the Exchange Rate Mechanism (ERM). Finally, as a result of the economic and political integration within the EU, Finland joined the Economic and Monetary Union (EMU) in 1999 and Euro replaced FIM in the financial market. Sweden, on the other hand, opted out from the EMU keeping Swedish currency floating against the Euro.

In addition to currency issues, both countries are interesting for their development economically. Originally, both countries had relatively closed financial markets which started to open up to foreign investors in the 1980s. Historically, Sweden was more developed economically and it had closer ties to the global financial markets. Therefore the development began earlier than in Finland. In Sweden, the regulation took mostly place in the 1980s. Final steps were taken in the beginning of 1990, when restrictions on foreign ownership were abolished. In Finland, the regulation started in the 1980s and ended in the beginning of the 1990s. At the beginning of 1993, all restrictions on foreign ownership were abolished.

3.2 Variables

We employ two types of risk factors in our international asset pricing model to represent economic risks, namely stock market risk and currency risk. The former is further decomposed into global stock market risk and local stock market risk as suggested by asset pricing models for mildly segmented stock markets. Global market portfolio returns are proxied by the return on the MSCI global equity market index with reinvested gross dividends. Local market portfolio returns are calculated from local market indices. Our second source of risk is related to exchange rate changes. As a proxy for the exchange rate risk, one can use either a global (trade-weighted) currency index or a single bilateral currency exchange rate. In this paper we choose the latter approach in order to detect if the USD/FIM or USD/SEK exchange rates are relevant for the pricing of Finnish or Swedish stocks, respectively. We use the continuously compounded change in the U.S. dollar value of FIM or SEK as measures of the country specific currency risk.

We test the model using three test assets in addition to the global market portfolio, namely the U.S., Finnish, and Swedish market portfolios. The U.S. stock market returns are calculated from the MSCI US total return index. The Finnish stock market returns from 1991 forward are calculated using the value-weighted Nasdaq OMXH yield index calculated by the stock exchange (previously named HEX index and covering all stocks quoted on the Main List). Prior to 1991, we use the WI-index which is calculated at the Hanken School of Economics.[6] For Sweden we use a similar index.[7]

Panel A of Table 2 shows summary statistics for the return series. Means and standard deviations are scaled by 12 and the square root of 12 to show them in annual terms. The annualized mean returns in USD for the world equity market and the US market are 9.081% and 9.024%, respectively. Similarly, the corresponding returns for Finland and Sweden are 13.614% and 12.911% per annum. Hence, Finland has offered the highest returns for US investors during the sample period, but in general both Sweden and Finland have offered more than two-times the excess return of the US market.

Similarly, the world and the US market portfolios show lower standard deviations as suggested by their lower returns. Finnish and Swedish stock markets share much higher volatilities. The Jarque-Bera test statistic indicates that all return series are non-normal. All stock markets but the USA show evidence of first-order autocorrelation. The autocorrelation is also surprisingly persistent in Finland and to smaller degree for the world stock portfolio as shown by the significant Ljung-Box $Q(12)$ test statistic. All series but the USDSEK exhibit high second moment dependencies as shown by the significance of the $Q^2(12)$ statistics.

To track predictable time-variation in asset returns, risk exposures, and the common rewards to risks, we use global and local predetermined forecasting variables. The variables are chosen on the basis of parsimony, previous empirical studies, and theoretical content (see, e.g., Ferson and Harvey, 1993; Harvey, 1995; De Santis and Gérard, 1998). The global information set contains: (1) a constant, (2) the global stock market return (LWRET), (3) the global stock market dividend

6 The WI-index has been frequently used to augment the HEX-index in previous studies. They are basically calculated similarly except for few minor differences. For more details on the indices, see Nyberg and Vaihekoski (2010).

7 The index was provided by Björn Hansson. More details on the Swedish index series can be seen from Frennberg and Hansson (1992).

yield in excess of one-month Eurodollar rate (XDYD), and (4) the U.S. default premium (USDP). All information variables are lagged by one period in order to be investors' conditioning information set.

LWRET is simply the lagged world stock market return calculated from the MSCI index. XDYD is calculated similar to De Santis and Gérard (1998), i.e., the return on the total return (gross) world MSCI index minus the return on the price index.[8] To get the excess dividend yield, we deduct the risk-free rate. USDP is the U.S. default premium measured as the difference in Moody's Baa minus Aaa bond yields.

When modeling the price of currency risk, we select two currency specific information variables for both currencies on top of the floating indicator variable (DFLO). The first variable is the difference between the Finnish (Swedish) and the U.S. one month interest rates (dINT). It is aimed at detecting devaluation risk in the short run as central banks typically increase the local interest rates to fight against the pressure of devaluation. Further, it is expected to capture longer-term pressure on the value of the Finnish (Swedish) currency. In practice, dINT was measured as the difference between the Finnish (Swedish) one month money market rate and the Eurodollar one month rate.[9] The second variable is the absolute value of lagged cross-currency return (|CCRET|), i.e., the lagged Swedish currency returns for Finnish currency risk, and vice versa. It is expected to capture devaluation risk and currency shocks in the short run and potential uncertainty in the long run in the other currency.

Finally, we use two variables to model changes in the price of local risk in the case of Sweden and Finland. The first is the same variables, dINT, as before. The second is a liberalization indicator, DLIB, which gets a value of one after 1990 for Sweden and 1993 for Finland when all restrictions on foreign ownership in the Swedish (Finnish) stock market were removed. Antell and Vaihekoski (2007) find the liberalization indicator to be a significant explanatory variable for the price of local risk in Finland.

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8 Note that this is not the same way MSCI calculates the dividend yield. Note also that from the beginning of year 2001, MSCI changed their index methodology which affected how dividends are taken into account in the gross indices. As a result, we calculate the dividend yield as 1/12 of the difference in one-year returns for the two indices.

9 We use money market rates are Helibor (1987-1998) and Euribor (1999-) for Finland and Stibor (1980-) for Swedish. Note that the money markets developed rather late in both countries. Prior to the availability of the money market rates, the Central Bank's base rate is used.

Panel B of Table 2 shows descriptive statistics for the information variables. None of the forecasting variables appear to be normally distributed and there is evidence of serial correlation. As a result, we test for the stationarity of the variables. Results (*p*-values) from the augmented Dickey-Fuller test where the test equation included a constant and a trend are reported. The number of lags is based on the Schwartz information criterion. Except for XDYD and USDP, the null hypothesis of a unit root can be rejected at the five percent level for all but one variable (dINT,FIN with a *p*-value of 7.5 per cent). Hence, we use the first differences of XDYD and USDP (denoted DXDYD and DUSDP) in the subsequent analysis.

Table 3 shows the correlation matrices for the test assets (Panel A) and for the conditioning instrumental variables (Panel B). The stock market returns are, as expected, correlated. The USA shows the highest correlation with the world (0.877), then Sweden (0.681) and Finland (0.541). The Finnish and Swedish currencies also have a high pairwise correlation (0.831). The instrumental variables show low correlations. The highest pairwise correlation is between the Finnish and Swedish dINT variables (0.769), but they are not used in the same equation. For most of the other variables, the correlation coefficients are below 0.2 in absolute terms. This suggests that none of the instrumental variables is likely to be redundant a priori.

4 EMPIRICAL RESULTS

4.1 Constant prices of risk

Our initial empirical tests concentrate on constant price of risk specifications of the asset pricing model with currency risk, outlined in equations (5)–(10). All prices of risk are assumed to be constant. We report results for a one-factor (global market risk), a two-factor (global and local market risk), and finally for a three-factor model (global, local, and currency risk). The results are reported in Panels A to C in Table 4. Diagnostic tests are provided in Panel D for the three factor model. For easier comparability, all six assets are included in all models.

Panel A in Table 4 shows the results for the global asset pricing model (ICAPM). The price of world risk is 0.023, but it is not found significant. This result is in line with previous research. In Panel B, the price of local market risk is not significant for the USA, again in line with previous studies. Interestingly, the local market risk for Finland and Sweden is highly significant. In Panel

C, we add the currency risk component into the model as a third risk factor. The price of currency risk is significant for both currencies together with the local market risk. The global market risk turns significant at the 10% level. The results in Panel C suggest that all three risk factors are relevant for the pricing of stocks in Finland and Sweden.

The variance process parameters (not reported) are all highly significant. Panel D reports some diagnostic tests for the three-factor model. The standardized residuals, defined as $z_t = \varepsilon_t / \sigma_t$, are theoretically mean zero with unit variance. The mean standardized residuals are fairly in line with the theoretical expected values (except for Sweden's –0.038). There is also some excess kurtosis left, rejecting the null hypothesis of normality. Despite a couple significant values for the test of autocorrelations in returns and squared returns, the residuals diagnostics are deemed acceptable.

4.2 Time-varying prices of global, local, and currency risk

Based on our results in Table 4, we continue with the three-factor model. Our full model allows prices of world, currency, and local risk to be time-varying, except for the USA, whose price of local market risk is assumed to be constant. The model is based on the model in Antell and Vaihekoski (2007). However, there are several modifications in addition to the ones mentioned in the introduction. First, the number of conditioning information has been reduced, and some of them have been replaced by new variables. These changes have been made to make the estimation more tractable. Second, to study the role of the fixed and floating regimes on the currency risk, our model allows the price of currency risk to differ before and after the floating decision in 1992. We do this by imposing a multiplicative specification for the price of currency risk using a country specific indicator variable for the period after the floating decision. In addition, we allow the price of risk to be a function of the cross-currency return shocks. Ultimately, our model for the price of currency risk is as follows:

$$\lambda_{t+1}^{c} = \kappa_0^c + \kappa_1^c DFLO^c + \kappa_2^c dINT_t^c + \kappa_3^c (dINT_t^c \times DFLO^c) + \kappa_4^c (\left|CCRET_t^c\right|) . \qquad (16)$$

The results for the price of risk parameters are reported in Panel A of Table 5. Note that the information variables have been demeaned (except DFLO) for the analysis to facilitate the interpretation of the results. In particular, the constant can be interpreted as the unconditional, long-term average. Panel B reports the variance process parameters. Panel C reports diagnostic tests. Finally, Panel D reports several Wald tests on the prices of risks.

For the world and local market risk the results are basically unchanged. The price of global market risk is significantly different from zero (the *p*-value from the Wald test being 0.011) and time-varying (*p*-value 0.047). The unconditional prices of local market risk for Finland and Sweden are significant. We also find the prices of local risk to be time-varying in Finland but not for Sweden (the *p*-values are 0.003 and 0.250 for Finland and Sweden, respectively). Somewhat surprisingly, contrary to Antell and Vaihekoski (2007), the liberalization indicator is not significant for either country. Their finding might be related to the floating decision which almost coincides with the liberalization for Finland.

The null hypothesis of zero price of currency risk can be clearly rejected in both countries as well as the hypothesis of constant price of currency risk. Analyzing the individual coefficients shows that almost all of them are significant. The unconditional price of currency risk is negative and almost equal for both countries as well highly significant (*p*-values less than one percent). Moreover, there is a positive and significant level shift after the floating decision reducing the price of currency risk (towards zero).

Interest rate difference between the local and the US interest rates seems to be a realistic variable in predicting the price of currency risk. The interest rate differential is a significant predictor for both countries (highly significant for Finland), the effect being negative prior and positive after the floating decision, indicated by the multiplicative effect $dINT \times DFLO$. Somewhat surprisingly, the results give mixed evidence on the relevance of the cross-currency currency shocks on the price of currency risk. For Finland, it is not significant, but for Sweden it is marginally significant (*p*-value 0.069). This could be due to fact that the currency turmoil often originated from Finland and later spread to Sweden. The sign of the parameter is as expected, negative, indicating that the higher the shock the higher the price of currency risk (in absolute terms).

Panel B reports the results for the variance process parameters. They are all highly significant. Panel C of Table 5 shows some diagnostic test statistics for the standardized residuals, which are all slightly positive. Normality is again rejected. However, the use of the quasi-maximum likelihood technique at least to some degree alleviates the problem.

4.3 Econometric considerations and robustness checks

The estimation was conducted using a modified version of a Gauss program originally created by Bruno Gérard in 1999 after which it has been used by several researchers as stated before. During the estimation process we came across a number of issues that have not been thoroughly brought forward in earlier literature. First, the results are sensitive to the stationarity of the used information variables. Even though the estimated variance process is stationary by construction, one should bear in mind that multiplying this stationary process with a non-stationary one yields a non-stationary process. The price of world market risk would be non-stationary unless we had taken the first difference of the world excess dividend yield, and the US default premium. Previous studies have generally not done this.

Second, in many cases one cannot test the usual restriction on the alphas in the system as the results become unstable after adding constants into the mean equations. This is shown for example by Lanne and Saikkonen (2006), who note that one should exclude the intercept term from the mean equation if it is not implied by theory. They show that the power of tests of the risk-return relation is severely hurt by the inclusion of a theoretically unnecessary intercept (even if it is statistically significant), and that the risk-return parameter gets very unstable in different samples. We also observe this instability by testing the one-factor model with a subsample from January 1980 forward. The price of market risk (not reported) varies much more under the intercept specification. Third, as a result of the previous problem, the adjusted pseudo *R*-squares tend to be low. Thus one has to use other diagnostic tests to validate the model under investigation.

As a test of the robustness of our results, we run a number of additional tests. First, we allow for asymmetry in the GARCH process following earlier studies (see, e.g., Bekaert and Wu, 2000; Cappiello et al., 2006). As a result, equation (13) is replaced with the following:

$$H_{t+1} = H_0 \circ (ii' - aa' - bb' - 0.5\, ii'dd') + aa' \circ \boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}'_t + bb' \circ H_t + dd' \circ \boldsymbol{\eta}_t\boldsymbol{\eta}_t', \qquad (17)$$

where $\boldsymbol{\eta}_t = \min(0, \boldsymbol{\varepsilon}_t) = \mathrm{I}_t\, \boldsymbol{\varepsilon}_t$, and I_t is a 6×6 matrix of indicators with off-diagonal elements equal to zero, and diagonal elements equal to one if the corresponding chock is negative, (i.e., $\mathrm{diag}(\mathrm{I}_t)_i = 1$ if ε_{it} is negative), d contains the diagonal elements of $D = dd'$. The term $-0.5ii'dd'$ is a modification of Ding and Engle (2001), and rests on the assumption that the distribution of $\boldsymbol{\varepsilon}_t$ is symmetric. Since D is diagonal, only own asymmetry is accounted for, which is similar to the univariate specification in Glosten et al. (1993).

The estimation is done for the three-factor models in Table 4 (Panel C) and Table 5. The results (not reported) are in effect similar, as are the residual diagnostics. The volatility asymmetry terms are jointly significant. However, they are driven by only a few of the markets. Modeling asymmetry does not seem to change the results and the diagnostics in this model set-up.

Next, we test the model using a subsample from June 1973 forward. Its start matches that of De Santis and Gérard (1998) and the beginning of the peg against a currency index for both currencies. The results are basically similar to those in Table 5. All parameter signs and magnitudes stay at the same level although some of them loose some of their statistical significance.

Finally, we test the model where also the price of currency risk is allowed to be a function of the global conditioning variables (similar to Antell and Vaihekoski, 2007). The results are again in line with those in Table 5. The only relevant parameter is the lagged world return, which is significantly negative both for FIM and SEK. The residual diagnostics get worse compared to those in Table 5.

5 SUMMARY AND CONCLUSIONS

In this paper we study the pricing of currency risk as well as global and local market risk in three stock markets using monthly data from March 1970 to August 2009. The three stock markets selected as our sample countries are the USA and two Nordic countries, namely Sweden and Finland. The Finnish and Swedish stock markets and currencies offer interesting test laboratories for many aspects of international asset pricing models. The long sample period includes, for example, a gradual liberalization of the financial markets and the decision to float the currencies both in Finland and in Sweden. Many East-European new EU members and e.g. China are currently experiencing a similar development.

In our empirical specification we utilize the multivariate GARCH-M framework of De Santis and Gérard (1998), allowing a time-varying variance-covariance process. First, we estimate constant price of risk versions of the asset pricing model. Both global, local and currency risks are significant. Second, we re-estimate the model allowing for time-varying prices of risk. The results show that the price of world risk is time-varying. However, even though its unconditional mean is positive with reasonable values, it is insignificant, which is in line with De Santis and Gérard (1998).

The price of local market risk is not priced in the US market. However, the local risk is priced in the Finnish and Swedish markets. These findings are partly conflicting with De Santis and Gérard (1998) who found that the local risk was not priced in any of the major stock markets in their study. Finding the local market risk relevant for the pricing of Finnish and Swedish stocks gives further evidence that one should consider partially segmented asset pricing models for smaller stock markets. The results also show that the price of local market risk is time-varying.

The price of currency risk is significantly different from zero and time-varying. Our specification for the price of currency risk is found to work well. The price of currency risk is in absolute terms significantly lower after the floating decision. Also the role of the interest rate differential in explaining the price of currency risk is different after the floating. Finally, somewhat surprisingly, we find only the return shock of the Finnish currency to price the Swedish currency risk, but not vice versa.

In this study we assume that investors price bilateral currency risk, not multilateral currency risk. However, if investors diversify across countries, the multilateral risk could be more relevant for the pricing of stocks. In addition, it would be interesting to study the effect of devaluation risk in more details as well as the role of cross-currency shocks in the variance process. These questions are left for future study.

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Table 1. Regimes and major changes to the value of the Finnish and Swedish currencies (Finnish Markka / Euro; and Krona)

Panel A lists Finnish currency regimes from 1949 to 2009. Panel B shows major changes to the value of the Finnish currency from 1971 to present day.

	Period (Finland)	Period (Sweden)
Panel A: Currency regimes		
Bretton Woods: Currency pegged against the USD	1949-1971/8	1951/8-1971/8
Smithsonian Agreement (Finland unofficially, Sweden officially)	1971-1973	1971-1973
Sweden joins the currency snake of the European Community		1973/3-1977/8
Markka fixed against trade-weighted currency index with fluctuation range, unofficial	1973/6-1977/11	
Peg against trade-weighted currency index with fluctuation range, official	1977/11-1991/6	1977-1991/5
Peg against the ECU (European Currency Unit) with fixed rate	7.6.1991	17.5.1991
Currency let floating	8.9.1992	19.11.1992
FIM joins the ERM with fixed central rate 5.80661/5.85424	14.10.1996	
FIM joins the EMU; Euro replaces FIM on financial markets	1.1.1999	
Euro notes are taken into use and Euro fully replaces FIM.	1.1.2002	
Panel B: Major changes to the value of FIM and SEK		
Gradual devaluation of 7.1%	1971-1974	
Devaluation against gold 1 %, revaluation against the USD 7.5 %		21.12.1971
Devaluation against gold 5 %, revaluation against the USD 5.6 %		16.2.1976
Devaluation against the German Mark (DEM) 3 %		18.10.1976
Devaluation 5.7 % (FIM) and 6 % (SEK)	5.4.1977	4.4.1977
Devaluation 2.9 % (FIM) and 10 % (SEK)	1.9.1977	29.8.1977
Devaluation 7.4 %	17.2.1978	
Revaluation 1.5% within fluctuation range	5.8.1979	
Revaluation 2.0% within fluctuation range widened	21.9.1979	
Revaluation 2.0% within fluctuation range	25.3.1980	
Devaluation 10 %		14.9.1981
Adjustment (devaluation) of 3.8 % within fluctuation range (FIM) and devaluation 10 % (SEK)	6.10.1982	8.10.1982
Devaluation 5.7%, fluctuation range reduced	11.10.1982	
Adjustment (revaluation) of 1.0% within fluctuation range	27.3.1984	
Adjustment of 1.6% within fluctuation range	1986	
Fluctuation range widened	1989	
Revaluation 3.8 %; fluctuation range changed	17.3.1989	
FIM devaluated 12.3%	15.11.1991	

Source: Bank of Finland, Bank of Sweden, Jonung (2000).

Table 2. Descriptive statistics for asset returns and information variables.

Descriptive statistics for continuously compounded monthly returns, and information variables. The global market portfolio is proxied by the MSCI total return world index. The US market return is proxied by the MSCI US index. The Finnish return is proxied by the WI-index (1970-1990) and HEX/OMXH index (1991-2009). The Swedish stock market return is from Frennberg and Hansson (1992). USDFIM and USDSEK are the logarithmic difference in the USD value of one Finnish Markka or Swedish Krona. The risk-free rate is calculated from Ibbotson (2009). All returns are measured in USD. The mean and standard deviation in Panel A are annualized (multiplied by 12 and the square root of 12, respectively). The global information set contains: world equity index return (LWRET), the world dividend yield in excess of risk-free rate (DXDYD), and the U.S. default premium (DUSDP). The last two variables are differenced once. The local information set contains the difference in the Finnish (Swedish) and the U.S. short-term interest rates (dINT) and absolute values of USDFIM and USDSEK. All information variables are lagged by one month. The sample size is 474 monthly observations from March 1970 to August 2009. The *p*-value for the Jarque-Bera test statistic of the null hypothesis of normal distribution is provided in the table. Q(12) and $Q^2(12)$ are the Ljung-Box statistics for the returns (information variables), and squared returns, respectively. In Panel B, the *p*-value is reported for the augmented Dickey-Fuller test for the null hypothesis of stationarity.

	Mean (%)	Std. dev. (%)	Skewness	Excess Kurtosis	Normality (p-value)	Autocorrelation[a] ρ_1	ρ_2	ρ_3	ρ_{12}	Q(12)[b]	$Q^2(12)$[b]
Panel A. Asset return series.											
World market portfolio	9.081	15.081	-0.814	2.428	<0.001	0.142*	-0.026	0.059	-0.022	0.054	<0.001*
Risk-free rate	4.433	0.563	0.199	0.357	<0.001	0.962*	0.935	0.914	0.725	<0.001*	<0.001*
U.S.	9.024	15.675	-0.681	2.638	<0.001	0.075	-0.021	0.047	0.067	0.362	0.003*
Finland	13.614	24.067	-0.254	2.606	<0.001	0.221*	-0.014	0.095	0.060	<0.001*	<0.001*
Sweden	12.911	22.657	-0.595	1.851	<0.001	0.133*	-0.032	0.104	0.020	0.127	<0.001*
USDFIM	0.031	10.192	-0.570	2.462	<0.001	0.065	0.026	0.039	-0.012	0.112	<0.001*
USDSEK	-0.805	10.664	-0.897	3.739	<0.001	0.125*	0.032	0.059	-0.022	0.030*	0.182

Table 2. *Continued*.

	Mean (%)	Std. dev. (%)	Skewness	Excess Kurtosis	Normality (p-value)	Autocorrelation[a] ρ_1	ρ_2	ρ_3	ρ_{12}	Q(12)[b]	ADF[c)] (p-value)
Panel B. Information variables.											
LWRET	0.755	4.352	-0.814	2.432	<0.001	0.139*	-0.028	0.058	0.067	0.065	<0.001
DXDYD	0.001	0.055	0.851	8.385	<0.001	-0.057	-0.013	-0.035	0.050	<0.001	<0.001
DUSDP	0.001	0.137	1.354	13.674	<0.001	0.267*	0.123	10.132	-0.053	<0.001	<0.001
dINT, FIN	0.427	3.507	0.299	1.665	<0.001	0.956*	0.919*	0.886*	0.711*	<0.001	0.075
dINT, SWE	0.859	3.482	1.054	6.170	<0.001	0.917*	0.867*	0.823*	0.587*	<0.001	0.017
|USDFIM|	1.114	1.957	2.643	10.507	<0.001	0.563*	0.510*	0.444*	0.341*	<0.001	<0.001
|USDSEK|	0.999	1.719	2.167	4.967	<0.001	0.451*	0.541*	0.498*	0.358*	<0.001	<0.001

a) Autocorrelation coefficients significantly (5%) different from zero are marked with an asterisk (*).
b) The *p*-value for the Ljung and Box test statistic for the null that autocorrelation coefficients up to 12 lags are zero.
c) MacKinnon (1996) one-sided *p*-values.

Table 3. Correlation matrices for asset returns and information variables.

Panel A provides the correlation matrix for the monthly returns for four equity markets (USA, Finland, Sweden, and the World), two currencies (Finnish Markka and Swedish Krona), and for the risk-free asset. Panel B provides the correlation matrix for the monthly values of the information variables. See Table 2 for an explanation of the variables.

Panel A: Correlation matrix for the test assets.

	USA	Finland	Sweden	USDFIM	USDSEK	World	Risk-free
USA	1						
Finland	0.407	1					
Sweden	0.548	0.621	1				
USDFIM	0.026	0.301	0.307	1			
USDSEK	0.134	0.329	0.410	0.831	1		
World	0.877	0.541	0.681	0.245	0.335	1	
Risk-free	0.018	0.001	0.005	-0.048	-0.068	-0.006	1

Panel B: Correlation matrix for the information variables.

	LWRET	DXDYD	DUSDP	dINT,F	dINT,S	|USDSEK|	|USDFIM|
LWRET	1						
DXDYD	0.075	1					
DUSDP	-0.104	0.250	1				
dINT, FIN	0.062	0.073	-0.088	1			
dINT,SWE	0.117	0.089	-0.058	0.769	1		
|USDSEK|	-0.111	-0.018	0.105	0.032	0.114	1	
|USDFIM|	-0.071	-0.025	0.074	0.116	0.258	0.725	1

Table 4. Integrated and partially segmented APM model with constant prices of global, currency, and local risk.

Quasi-maximum likelihood estimates of constant prices of risk are reported for one, two, and three factor models. The variance process is assumed to follow a multivariate GARCH(1,1) process. Reported *t*-values in parenthesis are based on QML robust standard errors (Bollerslev and Wooldridge, 1992). Panel D provides diagnostic tests for the three-factor model in Panel C. The sample size is 474 monthly observations from March 1970 to August 2009. Coefficients significantly (10%, 5% or 1%) different from zero are marked with one, two, or three asterisks, respectively.

Model tested	World	USA	Finland	Sweden	FIM/USD	SEK/USD
Panel A: One-factor model						
Price of world market risk, λ_w	0.023					
	(1.492)					
Panel B: Two-factor model						
Price of world market risk, λ_w	0.024					
	(1.474)					
Price of local market risk, λ_l		-0.060	0.011**	0.014***		
		(-1.463)	(2.276)	(3.009)		
Panel C: Three-factor model						
Price of world market risk, λ_w	0.023*					
	(1.747)					
Price of local market risk, λ_l		-0.001	0.013***	0.019***		
		(-0.294)	(2.758)	(3.924)		
Price of currency risk, λ_{fx}					-0.039**	-0.049***
					(-2.394)	(-3.234)
Panel D: Diagnostic tests (3-factor model)						
Avg. standardized residual (*z*)	0.002	0.014	-0.017	-0.038	-0.021	-0.009
Standard deviation of *z*	1.03	0.99	1.01	1.00	1.00	1.03
Skewness of *z*	-0.79	-0.27	-0.62	-0.49	-0.80	-0.95
Excess kurtosis of *z*	2.54	1.59	1.25	2.25	3.36	2.57
JB-test for normality, *p*-value	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001
Q(12), *p*-value	0.290	<0.001	0.238	0.113	0.074	0.049
$Q^2(12)$, *p*-value	0.479	0.078	0.008	0.903	0.989	0.369
Likelihood function	-7157.847	Akaike	30.278	Schwartz	30.436	

Table 5. Conditional partially segmented APM model allowing for different price of currency risk before and after the currency floating decision.

Quasi-maximum likelihood estimates of the conditional international CAPM with time-varying prices of risks. The model assumes that the U.S., Sweden, and Finland are partially segmented. The price of global risk is conditional on global information variables. The price of local market risk is assumed to be constant for the U.S.A., and time-varying for Sweden and Finland. Price of currency risk is conditional on local information variables which allows for testing the effect of currency floating decision. The global information set contains: the world equity index return (LWRET), the world dividend yield in excess of risk-free rate (DXDYD), and the U.S. default premium (DUSDP). The last two variables are differenced once. The local information set contains the difference in the Finnish (Swedish) and the U.S. short-term interest rates (dINT) and absolute values of cross-currency return on USDFIM or USDSEK (|CCRet|). All information variables are lagged by one month. Reported *t*-values in parenthesis are based on QML robust standard errors (Bollerslev and Wooldridge, 1992). In Panel D the *p*-values are reported in brackets. The sample size is 474 monthly observations from March 1970 to August 2009. Coefficients significantly (10%, 5% or 1%) different from zero are marked with one, two, or three asterisks, respectively.

		Conditioning information variables Z_{t-1}							
		Global			*Local*				
	Constant	DXDYD	DUSDP	LWRet	dINT	×DFLO	DFLO	|CCRet|	DLIB
Panel A. Parameter estimates for the prices of risk									
Price of world risk, λ_w	0.015	0.397	0.021	0.005*					
	(1.167)	(1.487)	(0.900)	(1.748)					
Constant price of local risk, λ_l									
USA	-0.002								
	(-0.456)								
Time-varying price of local risk, λ_{lt}									
Finland	0.027***				-0.005***				-0.002
	(3.179)				(-2.799)				(-1.388)
Sweden	0.022**				-0.021*				-0.006
	(2.414)				(-1.877)				(-0.553)
Price of currency risk, λ_c									
Finland	-0.084***				-0.016***	0.025***	0.079***	-0.269	
	(-3.921)				(-4.010)	(3.303)	(2.987)	(-1.183)	
Sweden	-0.085***				-0.009**	0.014**	0.066**	-0.989*	
	(-4.314)				(-2.225)	(2.085)	(2.514)	(1.820)	

Table 5. *Continued.*

Panel B: GARCH-parameters						
	USA	Finland	Sweden	FIM/USD	SEK/USD	World
a_i	0.210***	0.200***	0.163***	0.372***	0.214***	0.186***
	(7.896)	(6.732)	(3.391)	(11.806)	(5.202)	(6.694)
b_i	0.969***	0.972***	0.975***	0.836***	0.897***	0.977***
	(99.99+)	(99.99+)	(99.99+)	(99.99+)	(99.99+)	(99.99+)

Panel C: Diagnostic tests						
	USA	Finland	Sweden	FIM/USD	SEK/USD	World
Avg. standardized residual (*z*)	0.034	0.026	0.024	0.026	0.018	0.026
Standard deviation of *z*	1.02	0.99	1.01	0.99	1.00	1.03
Skewness of *z*	-0.74	-0.22	-0.59	-0.28	-0.72	-0.91
Excess kurtosis of *z*	2.29	1.58	1.24	1.30	2.94	2.44
JB-test for normality, *p*-value	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001
Q(12), *p*-value	0.284	<0.001	0.378	0.001	0.061	0.217
$Q^2(12)$, *p*-value	0.322	0.058	0.007	0.93	0.987	0.282
Likelihood function	-7117.329	Akaike	30.170	Schwartz	30.460	

Panel D: Robust Wald-tests			
Zero price for world risk, $\chi^2(4)$		12.93**	[0.011]
Constant price of world risk, $\chi^2(3)$		7.93**	[0.047]
Constant price of local risk, $\chi^2(2)$	Finland	11.93**	[0.003]
	Sweden	2.77	[0.250]
Zero price of currency risk, $\chi^2(5)$	FIM	37.14***	[<0.001]
	SEK	22.62***	[<0.001]
Constant price of currency risk, $\chi^2(4)$	FIM	27.26***	[<0.001]
	SEK	11.97**	[0.018]

Aboa Centre for Economics (ACE) was founded in 1998 by the departments of economics at the Turku School of Economics, Åbo Akademi University and University of Turku. The aim of the Centre is to coordinate research and education related to economics in the three universities.

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