Pricing of Foreign Exchange Risk from a German Investor’s Perspective

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Abstract:

We investigate the pricing of exchange risk using two international asset pricing model (IAPM) specifications: one model with a single currency index beside the world market factor (CI-IAPM) and a second employing three exchange risk variables to mirror the Triad regions (“Triad”-IAPM). For the total period we find significant exchange risk premia for both models. Sub-period and moving-window results are mixed and indicate pronounced time-variation. However, contrary to the CI-IAPM, the “Triad”-IAPM provide evidence on systematic exchange risk in more recent time. Consequently, the “Triad”-IAPM helps to gain a more differentiated picture on the pricing of exchange risk compared to most IAPMs which use only a single currency index.

Keywords: Foreign exchange risk; International asset pricing; GMM

JEL Classification: F31; G12; G15; C5

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1 Introduction

Since the breakdown of the Bretton-Woods Agreement in 1973, flexible exchange rates have become the subject of many empirical studies. One of the issues being of interest to scholars in the international finance and risk management literature has been the measurement of the exchange rate exposure of individual firms. From the perspective of an investor, however, the exchange rate exposure may not be relevant. Following the predictions of modern portfolio theory, an investor is only concerned about systematic risk which remains after all diversification opportunities have been exploited. Thus, investors will be compensated only for taking non-diversifiable or systematic risk. In this study, we empirically test for the existence of systematic exchange risk factors in an international asset pricing context. In order to obtain a more differentiated picture on the pricing of exchange risk, we decompose a single global currency index, which is generally used in previous studies, into three exchange risk factors representing the Triad regions and compare the results to those achieved with the broad currency index. Finding priced exchange rate exposure would have direct implications for the firm’s cost-of-capital determination, investment analysis, and, especially, for exchange risk hedging strategies of portfolio managers and corporate financial managers. For example, if exchange rate exposure is not regarded as a systematic risk factor, investors will not favorably value a company’s efforts to manage currency risk (see, e.g., Jorion, 1991, pp363).

1 The first major study on firms’ exchange risk exposure was carried out by Jorion (1990); more recent studies are from He and Ng (1998), Miller and Reuer (1998), Bodnar and Wong (2000), Glaum et al. (2000), and Williamson (2001).
We investigate the pricing of foreign exchange risk within the framework of an international asset pricing model (IAPM). De Santis et al. (1998, p2) suggest this approach saying that “the practical relevance of currency risk can be appropriately measured only within the context of an international asset pricing model.” There are two major groups of utility-based IAPMs which have evolved in international asset pricing research: first, IAPMs assuming that purchasing power parity (PPP) holds but allowing capital markets to be segmented from each other (see Black, 1974, and Stulz, 1981b). The second group of IAPMs assumes fully integrated capital markets but taking explicitly deviations from PPP into account (see, e.g., Solnik, 1974, Sercu, 1980, Stulz, 1981a, and Adler and Dumas, 1983). IAPM accounting for PPP deviations provide the theoretical foundation of our study.

The majority of empirical tests analyzing the existence of exchange risk premia in an IAPM framework has been conducted from a U.S. investor’s perspective, the U.S.-Dollar being the pricing currency (e.g. Ferson and Harvey, 1993, 1994, 1999, Dumas and Solnik, 1995, De Santis and Gérard, 1998, and Vassalou, 2000). Additionally, a Japanese investor’s perspective is taken by Tai (1999), whereas Oertmann (1997) studies an IAPM taking a Swiss investor’s view. Finally, De Santis et al. (1998), Hardouvelis et al. (2000) and Carrieri (2001) analyze the pricing of exchange risk from a German investor’s perspective with a special focus on the influence of the introduction of the European Monetary Union (EMU) on international asset

2 The APT has been extended into an international setting by Ross and Walsh (1983), Solnik (1983), and Ikeda (1991).
pricing. In general, most empirical studies find significant exchange risk premia, although the estimated exchange risk premia are not stable over time. In this paper, we investigate the existence of exchange risk premia taking a German investor’s view. In the first of the two model specifications we use a single trade-weighted currency index as the only exchange risk factor beside the world market factor. Most empirical studies employ such a single currency index, since this is regarded as a convenient way to incorporate exchange risk in an empirical IAPM. However, a single index is only a very rough and restrictive approximation to capture exchange risk in line with the underlying theoretical IAPMs. Therefore, we estimate our second model with several exchange risk factors. We split the broad currency index into three exchange risk factors which mimic the three major economic areas in the world, the Triad regions. Consequently, the bilateral DM/U.S.-Dollar and DM/Yen exchange rates and a trade-weighted European currency index act as exchange risk factors. Additionally, we use a moving-window approach in order to provide some evidence on the time-variation of exchange risk premia.

This paper is structured as follows: In Section 2, we review the IAPM of Adler and Dumas (1983) which is the theoretical foundation for our estimation models, followed by a data

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3 A number of studies test the pricing of exchange risk within a national asset pricing approach, assuming internationally segmented markets. For example, the U.S. stock market has been investigated by Jorion (1991), Dukas et al. (1996), the Japanese stock by Choi et al. (1998) and Doukas et al. (1999). Prasad and Rajan (1995) and Choi and Rajan (1997) examine several national stock markets.
section. The estimation results for both models will be reported in Section 3. The last section contains concluding remarks.

2 Model Specification and Data

2.1 Theoretical basis

In this paper we take Adler and Dumas’ (1983) IAPM as the starting point to develop empirical estimation models. It is derived in a L+1 country framework with stochastic inflation rates in all countries and assumes fully integrated capital markets. More recent studies show that the integration of capital markets has increased over time (see, e.g., Alford and Folks, Jr., 1996, Heston et al., 1995). Bekaert (1995, p76) states that markets in industrial economies, at least from 1980 onwards, are relatively well integrated. Additionally, Adler and Dumas’ model incorporates PPP deviations since they assume inflation rates to be stochastic.

Capital markets are integrated if assets with identical risk profiles have the same price, independent of the locations in which these assets are traded. This is the case when no cross-border investment barriers exist and all investors have access to the same investment opportunity set worldwide; see, e.g., Stulz (1981b), Bekaert and Harvey (1995).

A further indication of an integrated capital market can be seen in the tremendous increase in cross-border capital flows since the 1980s. For the U.S., cross-border capital flows of stocks amounted to 14.7% of GDP in the early 1990s (1990-1994), whereas this figure has risen to 56.5% in 1998. The corresponding figures for Germany are 15.2% and 69.8% respectively; see BIS (2000).
Deviations from PPP arise when inflation rates of two countries expressed in one currency differ, and, consequently, constitute real exchange rate risk. Under these assumptions, investors from one country hold two portfolios: the world market portfolio which is universal for all investors worldwide and a country-specific hedge portfolio which minimizes the risk of their consumption flows. The resulting multi-beta pricing condition is (Adler and Dumas, 1983, equation 14):

\[
E(R_i) = R_f + \theta \text{cov}(R_i, R_{WM}) + \theta \sum_{l=1}^{L} \left( \frac{1}{\theta^l} - 1 \right) W^l \frac{\text{cov}(R_i, \pi^l)}{W}
\]

where

\[
W = \sum_{l=1}^{L} W^l
\]

\[
\frac{1}{\theta} = \frac{\sum_{l=1}^{L} W^l / \theta^l}{W}
\]

\(R_i\) is the nominal return of asset \(i\), \(R_f\) the nominal domestic risk-free rate, and \(R_{WM}\) the world market portfolio’s nominal return. All returns are expressed in one currency, the pricing currency, to make returns comparable for investors of a given country. \(\pi^l\) stands for country \(l\)’s inflation rate, \(W^l\) represents the wealth of country \(l\) as a part of the world’s total wealth \((W)\), \(\theta^l\) reflects the degree of relative risk aversion in country \(l\), and \(\theta\) the average relative risk aversion. Equation (1) implies that the expected rate of return of asset \(i\) consists of the risk-

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6 Empirical studies show that the PPP equilibrium relationship does not hold in the short run and that it is difficult to prove that it holds in the long run; see, e.g., Rogoff (1996). The existence of PPP deviations leads to a situation where investors in different countries assign different real returns to the same assets, because they use different inflation rates to deflate nominal returns.
free interest rate of the domestic bond, the world market risk premium, and \(L+1\) additional inflation risk premia, stemming from \(L\) foreign inflation rates translated into the pricing currency and the domestic inflation rate (\(L+1\)st risk factor).

Under the additional assumption that inflation rates of all \(L+1\) countries are non-stochastic or zero, we can use \(L\) exchange risk premia instead of \(L+1\) inflation risk premia (see, e.g., Adler and Dumas, 1983, Dumas and Solnik, 1995, De Santis and Gérard, 1989). In this case, which we call “the Solnik-Sercu special case,” investors consider nominal exchange rate changes reflecting real exchange risk. With regard to the low volatility of inflation rates compared to exchange rate changes, this supplementary assumption seems to be justifiable. After substituting inflation rates of all \(L\) foreign countries, \(\pi_l\), with exchange rate changes of \(L\) foreign currencies to the pricing currency, \(s^l\), and after some further re-arranging, the model can be written as follows:

\[
E(R_i) = R_f + \delta_{WM} \text{cov}(R_i, R_{WM}) + \sum_{j=1}^{L} \delta_j \text{cov}(R_i, s^j)
\]

with \(\delta_{WM} = \theta = \frac{1}{\sum_{j=1}^{L} W^j} \frac{1}{\theta^j}\) and \(\delta_j = \theta \left( \frac{1}{\theta^j} - 1 \right) \frac{W^j}{W}\) \hspace{1cm} (2)
In equation (2), $\delta_{WM}$ represents the reward to world market covariance risk, and $\delta_l$ stands for the reward to exchange rate covariance risk introduced by unexpected changes of exchange rates. In risk premia notation the pricing relationship can now be written as follows:

$$E(R_i) = R_f + \lambda_{WM} \beta_{i,WM} + \lambda_{FX1} \beta_{i,FX1} + \lambda_{FX2} \beta_{i,FX2} + \cdots + \lambda_{FXL} \beta_{i,FXL}$$  \hspace{1cm} (3)

In equation (3) $\lambda_{WM}$ measures the world market risk premium, or the price of one unit of systematic world market risk, and $\beta_{i,WM}$ represents asset $i$'s exposure to the world market risk factor. $\lambda_{FX1}$ to $\lambda_{FXL}$ are exchange risk premia related to the $L$ exchange risk factors, and $\beta_{i,FX1}$ to $\beta_{i,FXL}$ are the exchange risk exposures or exchange rate sensitivities of asset $i$. Equation (2) and its risk premia version, equation (3) may serve as the theoretical justification to include our pre-specified risk factors in the empirical estimation model.

2.2 Model specification

In order to empirically test the existence of exchange risk premia, a linear $K$-factor model, in line with Merton’s (1973) intertemporal CAPM (ICAPM) or Ross’ (1976) APT, is employed. The following multi-beta model decomposes the actual return of asset $i$ into three components: its expected rate of return, the exposure contributions of $K$ systematic risk factors, and an idiosyncratic white noise error term:

7 In this formulation, the exchange rate is quoted with country $l$’s currency as the denominator and the pricing currency as the nominator.

8 Risk premia can also be obtained by multiplying the reward to covariance risk with the variance of the $j$-th risk factor [$\lambda_j = \delta_j \text{var}(R_j)$].
The expected rate of return of asset \( i \), \( E(R_i) \), itself depends on \( K \) risk premia, \( \lambda_j, j = 1, \ldots, K \), weighted by their sensitivities (betas) towards these \( K \) risk factors, plus the risk-free interest rate, represented by \( \lambda_0 \):

\[
E(R_i) = \lambda_0 + \sum_{j=1}^{K} \lambda_j \beta_{i,j} \quad \forall \quad i = 1, \ldots, N
\]  

The requested risk premia can be estimated within a system of \( N \) representative cross sections has to obtain the risk premia. Since equation (4) requires the included risk factors to be unexpected, we generate factor innovations using a simple vector autoregressive model (VAR). The VAR for each of the \( K \) risk factor is specified as follows:

\[
F_{j,t}^* = c_{j0} + c_{j1} F_{j,t-1}^* + \cdots + c_{jK} F_{K,t-1}^* + F_{j,t}
\]  

The superscript * stands for the total change in the value of the risk factor and the resulting residuals of each \( j \)-th regression, \( F_{j,t} \), represents the unexpected changes or factor innovations in the \( j \)-th risk factor. By construction, these factor innovations have zero means and can be employed as model adequate risk factors. Utilizing a VAR model with one lag reduces the number of observations by one for the following estimation models.

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9 This procedure has also been employed by Oertmann (1997), Brown and Otsuki (1993), Doukas et al. (1999).
Based on equation (3) we can specify two estimation models. The first model is a two-factor model, in which we use the world market factor and a single currency index to capture the effect of $L$ bilateral exchange risk factors following Ross and Walsh (1983).\footnote{When constructing a currency index, the weighting scheme becomes a pivotal role. Index weights in line with Adler and Dumas (1983) require information on the relative risk aversion of each country, $\theta^l$, and the portions of each country’s wealth to total wealth, $W^l$. Since these data are not obtainable, proxies such as trade-weights or GDP-weights are used instead; see, e.g., O’Brien and Dolde (2000) on this point.} Henceforth, this model will be called the currency-index IAPM (CI-IAPM). In excess-return formulation \((r_{i,t} = R_{i,t} - R_{f,t})\), the CI-IAPM can be expressed in matrix notation as follows:

\[
\begin{pmatrix}
  r_{1,t} \\
  \vdots \\
  r_{N,t}
\end{pmatrix}
= \lambda_{WM} \begin{pmatrix}
  \beta_{1,WM} \\
  \vdots \\
  \beta_{N,WM}
\end{pmatrix}
+ \lambda_{CI} \begin{pmatrix}
  \beta_{1,CI} \\
  \vdots \\
  \beta_{N,CI}
\end{pmatrix}
+ \begin{pmatrix}
  F_{WM,t} \\
  \vdots \\
  F_{CI,t}
\end{pmatrix}
+ \begin{pmatrix}
  \epsilon_{1,t} \\
  \vdots \\
  \epsilon_{N,t}
\end{pmatrix}
\]  
(7)

This estimation model implies the following restrictions: firstly, $\lambda_0$ is restricted to be zero, assuming that the zero-beta portfolio return is identical to the applied risk-free rate. Secondly, the world market risk premium, $\lambda_{WM}$, and the exchange risk premium, $\lambda_{CI}$, are restricted to be identical for all cross sections.

In order to develop a more realistic version in line with Adler and Dumas’ IAPM we include three exchange rate factors instead of a single currency index in our second model. By doing so, we try to reduce the theoretical disadvantages of a single currency index, such as possible averaging out effects, the choice of the “right” weights and the implicit assumption that...
weights do not change over time. We employ three exchange risk factors, mapping the economic regions of the Triad (“Triad”-IAPM). For the NAFTA region and the U.S.-Dollar pegged currency regions, the bilateral DM/U.S.-Dollar exchange rate is used, the Asian area is captured by the bilateral DM/Yen exchange rate, and a trade-weighted European currency index to cover the European economic region without Germany. The “Triad”-IAPM implies that the world market risk premium and all three exchange risk premia have to be equal for all cross sections, and can be written as follows:

$$
\begin{pmatrix}
    r_{1,t} \\
    \vdots \\
    r_{N,t}
\end{pmatrix} =
\lambda_{WM} \begin{pmatrix}
    \beta_{1,WM} \\
    \vdots \\
    \beta_{N,WM}
\end{pmatrix} + \lambda_{US} \begin{pmatrix}
    \beta_{1,US} \\
    \vdots \\
    \beta_{N,US}
\end{pmatrix} + \lambda_{Y} \begin{pmatrix}
    \beta_{1,Y} \\
    \beta_{N,Y}
\end{pmatrix} + \lambda_{EU} \begin{pmatrix}
    \beta_{1,EU} \\
    \beta_{N,EU}
\end{pmatrix} +
\begin{pmatrix}
    F_{WM,t} \\
    F_{US,t} \\
    F_{Y,t} \\
    F_{EU,t}
\end{pmatrix} \begin{pmatrix}
    \varepsilon_{1,t} \\
    \vdots \\
    \varepsilon_{N,t}
\end{pmatrix}
$$

(8)

2.3 Data

Our time sample encompasses the period from January 1981 to December 1998 (introduction of the Euro). Furthermore, we divide the total period into two sub-periods, nine years each. In this paper we employ continuously compounded excess returns using end-of-month data. The one-month Euro-DM money market rate is applied as the risk-free interest rate in order to calculate excess returns. Since we take the German investor’s position, all returns are expressed in Deutschmark, the underlying pricing currency. We use the following seventeen national MSCI stock markets as cross sections: Australia, Austria, Belgium, Canada,
Denmark, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the United Kingdom, and the United States of America plus Hong Kong. Since we utilize exclusively stock market indices as cross sections in our system, we make the implicit assumption that stock markets are segmented from other asset markets, especially bond markets. This implies that only international stocks, represented by the seventeen MSCI indices, could be used to diversify exchange risk in an international setting. Descriptive statistics on all national MSCI indices are displayed in Table 1, Panel A.

For both estimation models the world market risk factor is approximated by the MSCI World Index, a value-weighted performance index, which represents more than 90% of the world market capitalization. U.S.-Dollar returns of the MSCI World Index and all national MSCI indices have been translated into DM returns using the end of month DM/U.S.-Dollar exchange rate. For the CI-IAPM, we employ a broad trade-weighted currency index (CI) similar to the DM value published by the Deutsche Bundesbank. This currency index is quoted as

11 All MSCI indices used in this study are performance indices, where one-twelfth of the annual reported dividend yield is fully reinvested on every month’s end; see MSCI (1998). More information on the construction of MSCI indices can be found on the website of Morgan Stanley Capital International (http://www.msci.com/method/method.html).

12 Since the Deutsche Bundesbank’s DM value time-series is calculated on monthly average exchange rates, we construct a currency index using bilateral end-of-month exchange rate returns of the eighteen most important German trading partners and apply the Deutsche
the DM price for one unit of the currency basket. Consequently, a rise in the currency index indicates an depreciation of the DM, *et vice versa*. For the “Triad”-IAPM, we utilize the DM/U.S.-Dollar and the DM/Yen bilateral exchange rates and a trade-weighted European currency index (EU-11-CI), including eleven European countries. The development of these three currency factors and the broad currency index (CI) from January 1981 to December 1998 is shown in Figure 1. The period under consideration is characterized by intensive fluctuations of both bilateral exchange rates, the U.S.-Dollar and the Japanese Yen. The European (EU-11-CI) and the broad currency index display a similar pattern indicating a continuous appreciation of the Deutschmark over time. Descriptive statistics for all risk factors are given in Table 1 Panel B, and the correlations in Panel C. Finally, the hypothesis of normality cannot be rejected for most time series used in this study (see Harvey, 1995, on this GMM test for normality).

![Insert Figure 1 here](image)

Bundesbank’s weighting scheme, where the weights consist of direct bilateral trade activities (exports plus imports) and indirect effects on third markets ("Drittmarkteffekte"), see Deutsche Bundesbank (1998). All bilateral exchange rates used in this study are obtained from the Deutsche Bundesbank.

13 Included countries are: Austria, Belgium, Denmark, France, Italy, the Netherlands, Norway, Spain, Sweden, Switzerland, and the United Kingdom. The trade weights for the European currency index are based on those published by the Deutsche Bundesbank (1998).
3 Empirical Results

In this section we present the estimation results for two-factor model (CI-IAPM) and the four-factor model (“Triad”-IAPM). Following Ferson and Harvey (1994) and Oertmann (1997), we estimate both models using Hansen’s (1982) Generalized Method of Moments (GMM) as it has several advantages for this application. First, since risk premia and factor sensitivities are estimated simultaneously, errors-in-variables problems can be avoided. Second, the GMM approach does not require normally distributed asset and factor returns, and the error terms may be both serially correlated and heteroskedastic, which is often the case for financial data. The only prerequisites for applying this technique are the data to be strictly stationary and ergodic. In order to test the goodness-of-fit of the underlying model, Hansen’s (1982) J-test on over-identifying restrictions is applied. The test statistic is \( \chi^2 \) distributed with degrees of freedom equal to the number of restrictions imposed by the model. In line with Ferson & Harvey (1994) and Oertmann (1997), a vector of ones and the contemporaneous changes of the pre-specified risk factors, \( F^*_{1,t}, \ldots, F^*_{K,t} \), are applied as instruments to generate moment conditions. The resulting orthogonality conditions require that \( E(\varepsilon_i, t | F^*_{j,t}) = 0 \) and \( E(\varepsilon_i, t) = 0 \), for all \( i = 1, \ldots, N \) and \( j = 1, \ldots, K \).\(^{14}\)

\(^{14}\) For the calculations we use an iterated GMM procedure, since this technique has superior small-sample properties over the two-stage GMM (see Ferson and Foerster, 1994). Moreover, the Newey-West’s (1987) heteroskedasticity and autocorrelation consistent covariance matrix is chosen as the weighting matrix of the GMM estimator.
3.1 The two-factor model: CI-IAPM

The estimation results for the CI-IAPM for the total period and both sub-periods are shown in Table 2. The GMM statistics cannot be rejected, indicating a reasonable data fit to the restrictions imposed by the CI-IAPM for all sample periods. For the total period we estimate a significant world market risk premium of 8.11% p.a. and a highly significant and negative exchange risk premium of -3.62% p.a. The negative currency risk premium deserves some deeper explanation. Referring to equation (2) a negative exchange risk premium in the CI-IAPM will arise, if the investor’s relative risk aversion, $\theta$, is greater than one.  

This can be shown when the third term on the right-hand side of equation (2) is substituted by a single currency risk measure. Under the assumption that the currency index reflects the true weights of all $L$ currencies, the following equation (9) proves that the reward to exchange rate covariance risk, $\delta_{CI}$, is equal to $(1-\theta)$ (see O’Brien and Dolde, 2000). With $\lambda_{CI} = \delta_{CI} \text{var}(F_{CI})$ and $\theta > 1$, a negative currency risk premium arises:

15 If an investor’s relative risk aversion equals one, the investor is called a log investor and holds only the world market portfolio and no exchange risk hedge portfolio. The more the risk aversion coefficient exceeds one, the more investors are willing to invest into their individual exchange risk hedge portfolio which results in an increased negative exchange risk premium. A relative risk aversion of two ($\theta=2$) is frequently accepted; see Adler and Dumas (1983, footnote 47).
\[ \delta_{ci} = \theta \sum_{i=1}^{L} \left( \frac{1}{\theta_i} - 1 \right) \frac{W_i}{W} = \theta \left( \sum_{i=1}^{L} \left( \frac{W_i}{W} \times \frac{1}{\theta_i} \right) - \sum_{i=1}^{L} \left( \frac{W_i}{W} \right) \right) = \theta \left( \frac{1}{\theta} - 1 \right) = (1 - \theta) \]  

(9)

In combination with the respective exchange risk beta, a negative exchange risk premium would mean that investors require a lower rate of return for an asset with a positive exchange risk beta, \textit{ceteris paribus}. Investors are willing to forgo a part of the required rate of return of these assets, since these assets serve as instruments to hedge exchange risk. In other words, investors are prepared to pay a premium on assets which offer an exchange risk hedging possibilities (see De Santis et al., 1998). As foreign assets generally provide good potential for hedging exchange risk, foreign stock markets (especially these outside the European Monetary System, EMS) are expected to show positive exchange risk betas from a German investor’s perspective.\[16\]

[Insert Table 2 here]

To get some information on the stability of exchange risk premia over time, the CI-IAPM is calculated for both sub-periods. For the first sub-period, we estimate a positive significant world market risk premium and a negative, highly significant currency risk premium of -11.18% p.a. For sub-period 2, however, we find an insignificant exchange risk premium of

\[ \text{16 Stock markets with good hedging potential, indicated by a significant positive exchange risk beta, are Australia, Canada, Hong Kong, Italy, USA. We find significant negative betas for Switzerland, France, and the Netherlands, whose currencies are strongly linked to the Deutschmark. As expected, the domestic German stock market displays the highest negative beta of -0.873. Estimation results are available from the author on request.} \]
-0.87% p.a. and a very low but significant world market risk premium of 1.83% p.a. Hence, our results indicate that exchange risk is perceived as a relevant pricing factor by the investors only in the 1980s, whereas in the 1990s exchange risk is not perceived as a systematic risk factor. The significant exchange risk premium from the first sub-period might be explained by the extreme movements of the DM/U.S.-Dollar and the DM/Yen exchange rate (see Figure 1). Besides, both currencies have a high weight in the currency index, the U.S.-Dollar accounts for 14.8% and the Japanese Yen for 12.3%. In sub-period 2, the U.S.-Dollar and the Japanese Yen do not exhibit such pronounced patterns. Furthermore, after the Maastricht Treaty in February 1992 it could be expected that exchange risk stemming from European currencies might decrease due to the enforcement of the convergence criteria to prepare for the single European currency. Since European currencies account for an aggregated index weight of 71.8%, an insignificant currency index risk premium in sub-period 2 does not seem to be surprising.

Finally, we apply a moving-window approach to display the time-variation of exchange risk premia more specifically. The underlying procedure can be explained as follows: in the first step, the CI-IAPM is estimated using data from January 1981 to December 1989. In the following step, the estimation window is moved one month forward and the model is estimated again. This is repeated 108 times until the last estimation period from January 1990 to De-

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17 Some interruptions of this process of “convergence” occurred in September 1992 following the German reunification, when Britain and Italy were forced to withdraw from the EMS, and in August 1993, when the French franc was under enormous pressure. For more detailed information on the history of the EMS, see, e.g., Levi (1996).
December 1998 is reached. Consequently, the resulting graph, depicted in Figure 2, shows the gradual evolution of exchange risk premia over time. The bars in Figure 2 specify the level of significance for each estimated exchange risk premium. In the first half of the 109 coefficients, the estimated risk premia are negative and, with few exceptions, highly significant. When we move the windows forward and include more and more data from the mid and late nineties, the exchange risk premia converge to zero and become insignificant. At this point the question arises whether we can conclude that in more recent time a German investor has no longer been exposed to systematic exchange risk.

[Insert Figure 2 here]

3.2 The four-factor model: “Triad”-IAPM

In order to gain a more differentiated picture on the existence of exchange risk premia, we split the single currency index factor into three exchange risk factors covering the economic regions of the Triad. This appears especially interesting, since we often observe opposite developments of the U.S.-Dollar and the Japanese Yen against the Deutschmark from 1986 onwards which could have been cancelled out within a currency index. Therefore, we expect to find more pronounced effects related to the U.S.-Dollar and the Japanese Yen when applied separately. On the contrary, we do not expect to find strong support for a priced European currency risk factor, since most of the included exchange rates in the European index are those of EMS currencies, which show relatively low fluctuations over time, especially in more recent time (in the 1990s) when EMU members prepared for the Euro.

In Table 3 we present estimation results for the total period and both sub-periods. The GMM tests for the total period and both sub-periods cannot be rejected and its p-values are considerably higher compared to those of the CI-IAPM. For the total period, significant risk premia
can be found for all included risk factors. The world market risk premium is positive with 8.98% p.a.; all three exchange risk premia are negative. We estimate highly significant and negative risk premia for the U.S.-Dollar and the Japanese Yen (-15.42% p. a. resp. -22.69% p.a.), the EU currency risk premium is smaller in size (-2.26% p.a.) and significant at the 5% level. For both sub-periods, we estimate positive and highly significant risk premia for the world market risk factor, similar to the total period’s result. As far as exchange risk premia are concerned, a look at both sub-periods reveals a surprising result. For the first sub-period, we calculate a highly significant European currency risk premium of -2.39% p.a. and a U.S.-Dollar risk premium of -6.57% p.a., being significant at the 10% level. Moreover, we find a positive Yen exchange risk premium of 4.16% p.a., which is not significant. However, for the second sub-period, we obtain a highly negative and significant Yen exchange risk premium of -28.98% p.a. and a U.S.-Dollar exchange risk premium of -6.08% p.a. (10% level). The European currency risk premium is close to zero (0.66% p.a.) and insignificant.

At this point, it can be summarized that the exchange risk premia vary considerably between both sub-periods. This is especially evident for the Yen and the European risk factor, where we find coefficients of different signs and magnitudes next to changing significance. These findings support the former conjecture that in the advent of the EMU German investors do not consider European currency risk as a systematic risk factor. Moreover, as expected, the

18 De Santis et al. (1998) investigated the impact of European currency risk on international asset pricing. They found that EMU currency risk is of little relevance compared to non-EMU currency risk and conclude that the introduction of the Euro in January 1999 is unlikely to have a large effect on international asset pricing.
U.S.-Dollar risk factor is priced in both sub-periods. Surprisingly, we do not find a significant exchange risk factor for the Japanese Yen in the first sub-period, however. To conclude, applying the “Triad”-IAPM we find evidence of priced exchange risk factors for the more recent period which we do not when using a single currency index.

We also run the moving-window procedure with the “Triad”-IAPM. The resulting graphs for all three exchange risk premia are given in Figure 3. The first chart exhibits the U.S.-Dollar risk premia over time. It is noticeable that this risk premium displays a large bandwidth, ranging from -21.69% p.a. to +14.16%. We observe that after some insignificant U.S.-Dollar risk premia, we get significant negative risk premia for the 1980s. As we move the window forward, estimated risk premia become positive and with data from 1985 onwards even significantly different from zero. However, the sign switches back to negative in the 1990s, predominantly with significant coefficients. In sum, this graph shows that from a German investor’s perspective, exchange risk introduced by DM/U.S.-Dollar exchange rate is perceived differently in different time periods. Furthermore, we find a conspicuous pattern of significant negative and positive exchange risk premia over time. The second chart depicts the Yen exchange risk premium, which is also not stable over time and varies even stronger than the U.S.-Dollar risk premium (from -61.24% to +25.64% p.a.). Surprisingly, the estimated risk premia show an almost identical pattern to that of the U.S.-Dollar risk premium, but more pronounced in the second half. Thus, we find that Yen exchange risk is generally perceived as a systematic risk factor from the perspective of German investors, except for some periods in the first half of the estimations. Although the degree of a country’s risk aversion may provide some explanations for the changing signs of exchange risk premia, it is hardly comprehensible that the degree of risk aversion changes that dramatically to explain the pronounced shifts of
the U.S.-Dollar and the Japanese Yen risk premia in Figure 3. These striking findings require further explanations which cannot supplied yet. The risk premium of the European currency index, in contrast, is mainly negative and close to zero, and significant risk premia are more the exception than the rule. We find few significant risk premia only in the first half, whereas in the second half, practically all estimated risk premia showed up to be insignificant. These findings could be interpreted as the result of the efforts of the members of the EMU in the advent of the Euro.

4 Conclusion

The objective of this study is to investigate whether exchange risk is priced in an international stock market. Under the assumption that equity markets are integrated internationally, we test two empirical IAPMs. In the first model (CI-IAPM) we employ a single trade-weighted currency index beside the world market portfolio, and for the second model (“Triad”-IAPM), we split the global currency index into three exchange risk factors mapping the Triad regions.

Applying the CI-IAPM, we obtain significant negative exchange risk premia for the total period and mixed results for both sub-periods. Moving-window results confirm this time-variation of the currency index risk premium and show that especially in more recent time-periods the estimates are not statistically significant. In order to investigate the existence of exchange risk premia in more detail, we estimate the “Triad”-IAPM and find significant negative risk premia for all three exchange risk factors. However, the Yen and the European exchange risk premia vary considerably between both sub-periods, whereas U.S.-Dollar risk practically does not change. The finding of significant exchange risk premia for the U.S.-Dollar and the Japanese Yen in the more recent time period, in comparison to the CI-IAPM, leads us to conclude that using a global currency index as the only exchange risk factor might produce deceptive results on the existence of exchange risk premia. Therefore, a major con-
tribution of this study lies in the economical decomposition of the single trade-weighted currency index into a smaller number of exchange risk factors which helps to gain a more differentiated picture on the pricing of exchange risk. Results from a moving-window procedure show that both bilateral exchange risk factors exhibit the same distinctive pattern over time with high risk premia in absolute terms and changing signs.
References


Table 1: Descriptive statistics for all cross sections and risk factors (January 1981 to December 1998)

### PANEL A

<table>
<thead>
<tr>
<th>Stock markets</th>
<th>Mean (in % p.a.)</th>
<th>S.D. (in % p.a.)</th>
<th>Skewness</th>
<th>Kurtosis</th>
<th>GMM test of normality</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>1.484</td>
<td>28.882</td>
<td>-2.303</td>
<td>18.791</td>
<td>5.490 *</td>
</tr>
<tr>
<td>Austria</td>
<td>2.508</td>
<td>23.427</td>
<td>0.071</td>
<td>5.607</td>
<td>4.218</td>
</tr>
<tr>
<td>Belgium</td>
<td>12.656</td>
<td>18.639</td>
<td>-0.210</td>
<td>6.998</td>
<td>2.654</td>
</tr>
<tr>
<td>Canada</td>
<td>0.850</td>
<td>22.154</td>
<td>-0.708</td>
<td>5.838</td>
<td>5.069 *</td>
</tr>
<tr>
<td>Denmark</td>
<td>8.043</td>
<td>18.855</td>
<td>-0.338</td>
<td>3.204</td>
<td>5.518 *</td>
</tr>
<tr>
<td>France</td>
<td>7.783</td>
<td>21.452</td>
<td>-0.747</td>
<td>4.809</td>
<td>5.165 *</td>
</tr>
<tr>
<td>Germany</td>
<td>7.972</td>
<td>20.498</td>
<td>-0.918</td>
<td>5.575</td>
<td>3.097</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>4.938</td>
<td>37.156</td>
<td>-1.265</td>
<td>9.059</td>
<td>1.987</td>
</tr>
<tr>
<td>Italy</td>
<td>5.690</td>
<td>27.792</td>
<td>0.606</td>
<td>3.353</td>
<td>1.643</td>
</tr>
<tr>
<td>Japan</td>
<td>2.529</td>
<td>24.328</td>
<td>-0.133</td>
<td>3.491</td>
<td>3.097</td>
</tr>
<tr>
<td>Netherlands</td>
<td>11.777</td>
<td>17.817</td>
<td>-0.834</td>
<td>6.253</td>
<td>2.700</td>
</tr>
<tr>
<td>Norway</td>
<td>2.076</td>
<td>27.478</td>
<td>-0.992</td>
<td>6.331</td>
<td>2.700</td>
</tr>
<tr>
<td>Spain</td>
<td>10.640</td>
<td>24.558</td>
<td>-0.509</td>
<td>4.848</td>
<td>3.515</td>
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<tr>
<td>Sweden</td>
<td>12.696</td>
<td>25.409</td>
<td>-0.567</td>
<td>3.994</td>
<td>3.733</td>
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<tr>
<td>Switzerland</td>
<td>9.221</td>
<td>18.078</td>
<td>-0.883</td>
<td>6.188</td>
<td>2.411</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>7.659</td>
<td>20.266</td>
<td>-1.089</td>
<td>7.335</td>
<td>2.562</td>
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<tr>
<td>United States</td>
<td>8.928</td>
<td>19.627</td>
<td>-0.958</td>
<td>6.396</td>
<td>4.515</td>
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### PANEL B

<table>
<thead>
<tr>
<th>Risk factors</th>
<th>Mean (in % p.a.)</th>
<th>S.D. (in % p.a.)</th>
<th>Skewness</th>
<th>Kurtosis</th>
<th>GMM test of normality</th>
</tr>
</thead>
<tbody>
<tr>
<td>MSCI World (WM)</td>
<td>6.812</td>
<td>16.858</td>
<td>-1.001</td>
<td>6.071</td>
<td>4.835 *</td>
</tr>
<tr>
<td>Trade-weighted currency index (CI)</td>
<td>-1.462</td>
<td>3.799</td>
<td>-0.002</td>
<td>3.176</td>
<td>0.320</td>
</tr>
<tr>
<td>DM/U.S.-Dollar (US$)</td>
<td>-0.877</td>
<td>11.443</td>
<td>0.121</td>
<td>3.439</td>
<td>0.577</td>
</tr>
<tr>
<td>DM/Yen (Yen)</td>
<td>2.244</td>
<td>10.458</td>
<td>0.594</td>
<td>4.278</td>
<td>4.471</td>
</tr>
<tr>
<td>European CI (EU-11-CI)</td>
<td>-2.089</td>
<td>2.693</td>
<td>-0.885</td>
<td>5.495</td>
<td>5.786 *</td>
</tr>
</tbody>
</table>

### PANEL C

<table>
<thead>
<tr>
<th></th>
<th>WM</th>
<th>CI</th>
<th>US$</th>
<th>Yen</th>
<th>EU-11-CI</th>
</tr>
</thead>
<tbody>
<tr>
<td>WM</td>
<td>1.000</td>
<td>0.578 ***</td>
<td>0.529 ***</td>
<td>0.413 ***</td>
<td>0.365 ***</td>
</tr>
<tr>
<td>CI</td>
<td>1.000</td>
<td>0.815 ***</td>
<td>0.647 ***</td>
<td>0.752 ***</td>
<td></td>
</tr>
<tr>
<td>US$</td>
<td></td>
<td>1.000</td>
<td>0.358 ***</td>
<td>0.437 ***</td>
<td></td>
</tr>
<tr>
<td>Yen</td>
<td></td>
<td></td>
<td>1.000</td>
<td>0.186 ***</td>
<td></td>
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<tr>
<td>EU-11-CI</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>1.000</td>
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</tbody>
</table>

In Panel A and B, sample means and standard deviations for excess DM returns are given in % p. a. A GMM test for normality is conducted, testing the hypothesis that the coefficient for skewness and excess kurtosis are jointly equal to zero. The goodness-of-fit test for the model is $\chi^2$ distributed with two degrees of freedom. * / ** / *** symbolize that the null hypothesis of normally distributed returns is rejected on the 10% / 5% / 1% level of significance. Panel C provides the correlation matrix of the actual risk factor returns (no factor innovations).
Table 2: Risk premium estimates for CI-IAPM

<table>
<thead>
<tr>
<th></th>
<th>World market risk premium $\lambda_{WM}$</th>
<th>Exchange risk premium $\lambda_{CI}$</th>
<th>GMM $\chi^2$ statistic p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Total period</strong></td>
<td>8.112 ***</td>
<td>-3.615 ***</td>
<td>18.719</td>
</tr>
<tr>
<td>(81:1-98:12)</td>
<td>11.664</td>
<td>-3.456</td>
<td>0.227</td>
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<tr>
<td><strong>Sub-period 1</strong></td>
<td>12.471 ***</td>
<td>-11.182 ***</td>
<td>12.299</td>
</tr>
<tr>
<td>(81:1-89:12)</td>
<td>7.461</td>
<td>-4.812</td>
<td>0.656</td>
</tr>
<tr>
<td><strong>Sub-period 2</strong></td>
<td>1.827 ***</td>
<td>-0.873</td>
<td>15.135</td>
</tr>
<tr>
<td>(90:1-98:12)</td>
<td>2.802</td>
<td>-1.078</td>
<td>0.442</td>
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</table>

The reported risk premia are estimated using GMM and shown in % p.a. The $t$-statistics reported underneath are heteroskedasticity and serial correlation consistent estimates. The GMM test statistics for goodness-of-fit are heteroskedasticity and serial correlation consistent and distributed with 15 degrees of freedom; p-values are reported below. * / ** / *** denotes that coefficients are significant at the 10% / 5% / 1% level.
### Table 3: Risk premium estimates for “Triad”-IAPM

<table>
<thead>
<tr>
<th></th>
<th>World market risk premium $\lambda_{WM}$</th>
<th>U.S.-Dollar risk premium $\lambda_{US}$</th>
<th>Yen risk premium $\lambda_{Y}$</th>
<th>EU risk premium $\lambda_{EU}$</th>
<th>GMM $\chi^2$ statistic</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Total period</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>8.530</td>
<td>-3.697</td>
<td>-4.866</td>
<td>-2.370</td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Sub-period 1</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(81:1-89:12)</td>
<td>7.395 ***</td>
<td>-6.574 *</td>
<td>4.160</td>
<td>-2.390 ***</td>
<td>10.201</td>
<td>0.677</td>
</tr>
<tr>
<td></td>
<td>6.883</td>
<td>-1.685</td>
<td>1.113</td>
<td>-2.587</td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Sub-period 2</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(90:1-98:12)</td>
<td>6.304 ***</td>
<td>-6.081 *</td>
<td>-35.693 ***</td>
<td>0.658</td>
<td>9.527</td>
<td>0.732</td>
</tr>
<tr>
<td></td>
<td>4.467</td>
<td>-1.914</td>
<td>-6.739</td>
<td>0.754</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

The reported risk premia are estimated using GMM and shown in % p.a. The $t$-statistics reported underneath are heteroskedasticity and serial correlation consistent estimates. The GMM test statistics for goodness-of-fit are heteroskedasticity and serial correlation consistent and distributed with 13 degrees of freedom; p-values are reported below. *, **, *** denotes that coefficients are significant at the 10%, 5%, 1% level.
Figure 1: Exchange risk factors over time (January 1981 – December 1998)

Source: Deutsche Bundesbank, own calculations (January 1981 = 100).
The graph depicts 109 exchange risk premia in % p.a., employing the CI-IAPM. An estimation window of nine years is used to calculate these exchange risk premia, applying an iterative GMM procedure. This window is moved further by one-month steps. The shaded bars represent the p-values of the corresponding parameter test of significance.

Figure 2: Exchange risk premium over time (CI-IAPM)
Figure 3: Exchange risk premia over time (“Triad”-IAPM)

U.S.-Dollar Exchange Risk Premium over Time
(Cross sections: 17 MSCI indices, Sample size: 9 years, Step size: 1 month)

Yen Exchange Risk Premium over Time
(Cross sections: 17 MSCI indices, Sample size: 9 years, Step size: 1 month)

EU Currency Index Exchange Risk Premium over Time
(Cross sections: 17 MSCI indices, Sample size: 9 years, Step size: 1 month)