Asymmetric Currency Exposure and Currency Risk Pricing

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Abstract

Whether stock returns are linked to currency movements and whether currency risk is priced in a domestic context are less conclusive and thus still subject to a great debate. Based on a different approach, this paper attempts to provide new empirical evidence on these two inter-related issues, which are critical to investors and corporate risk management. In particular, this paper not only explores the possibility of asymmetric currency exposure that may explain why prior studies, which focus exclusively on linear exposure, have difficulty in detecting it, but also tests whether this asymmetric currency exposure is priced. The result shows that more than 50% of US industries are asymmetrically affected by currency movements based on the tests of conditional asset pricing model with multivariate GARCH in mean (MGARCH-M) parameterization where both conditional first and second moments of asset returns and risk factors are estimated simultaneously. Moreover, when the model utilizes individual security returns, 90% of the cases are asymmetric. Finally, the currency risk is priced regardless of whether bilateral or multilateral exchange rate is used, whether returns are measured at weekly or monthly horizon, and whether aggregate industry portfolios or individual securities are used to estimate the model. The strong evidence of asymmetric currency exposure and currency risk pricing suggests that both asymmetry and conditional heteroskedasticity play important roles in testing currency exposure and its price.

JEL Classifications: C32, G12

Key Words: Time-varying Currency Risk premium; Asymmetric Currency exposure; Multivariate GARCH-M

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Whether stock returns are linked to currency movements and whether currency risk is priced in a domestic context are less conclusive and thus still subject to a great debate. Based on a different approach, this paper attempts to provide new empirical evidence on these two inter-related issues, which are critical to investors and corporate risk management. In particular, this paper not only explores the possibility of asymmetric currency exposure that may explain why prior studies, which focus exclusively on linear exposure, have difficulty in detecting it, but also tests whether this asymmetric currency exposure is priced. The result shows that more than 50% of US industries are asymmetrically affected by currency movements based on the tests of conditional asset pricing model with multivariate GARCH in mean (MGARCH-M) parameterization where both conditional first and second moments of asset returns and risk factors are estimated simultaneously. Moreover, when the model utilizes individual security returns, 90% of the cases are asymmetric. Finally, the currency risk is priced regardless of whether bilateral or multilateral exchange rate is used, whether returns are measured at weekly or monthly horizon, and whether aggregate industry portfolios or individual securities are used to estimate the model. The strong evidence of asymmetric currency exposure and currency risk pricing suggests that both asymmetry and conditional heteroskedasticity play important roles in testing currency exposure and its price.

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I. Introduction

Since the breakdown of the fixed exchange system in 1973, currency variability has been a subject of interest and concern. Because of the increasing globalization of product and financial markets, currency movements have become an important source of risk for a firm operating in an international environment. Under this environment, the firm will be interested in knowing first whether its operating cash flows are exposed to currency movements, and if they are, then the firm should concern whether to hedge the currency risk. To answer these two interrelated issues – currency exposure (or exchange rate exposure) and currency risk pricing, the firm must establish the facts that not only are its operating cash flows significantly affected by the currency movements, but also the currency risk is priced in the asset market in equilibrium in order to justify its currency hedging. According to standard portfolio theory if the effect of currency risk does not vanish in well-diversified portfolios, exposure to this risk should command a risk premium in the sense that investors are willing to pay a premium to avoid this systematic risk. In this case, hedging policies can affect the cost of capital of a firm and the firm who actively engages in currency hedging is justifiable.\(^1\) On the other hand, if currency risk is diversifiable, investors are not willing to pay a premium for firms with active hedging policies since investors can diversify the risk themselves (see, e.g., Dufey and Srinivasulu (1983), Smith and Stulz (1985), Jorion (1991)). Consequently, if

\(^1\) Since the primary goal of a firm is not to reduce its cost of capital, when I state that the firm’s cost of capital will be reduced through currency hedging if currency risk is priced, I implicitly assume that the firm will undertake the projects with positive NPVs. Otherwise, the firm can buy T-bills and lower its systematic risk and cost of capital, but this is zero NPV.
there is evidence of significant currency exposure, then whether it is one of the priced factors in the sense of Ross’s arbitrage pricing theory (APT) (1976) is a critical issue for both investors and corporate financial managers.

Surprisingly many studies have examined above two issues in isolation. According to Adler and Dumas (1980, 1984), currency exposure can be obtained by regressing changes in firm value against exchange rate changes. However, previous empirical studies have very limited success in detecting it (e.g., Jorion (1990), Bodnar and Gentry (1993), Bartov and Bodnar (1994), Choi and Prasad (1995), Chow et al. (1997), He and Ng (1998), Allayannis and Ofek (2001), Griffin and Stulz (2001), among others). For example, Jorion (1990) finds that the currency exposure of 287 US multinationals is mostly insignificant. Subsequent studies by Bodnar and Gentry (1993), Bartov and Bodnar (1994), Choi and Prasad (1995) find similar results. Using Japanese data, He and Ng (1998) find that only 25% of the 171 multinationals in their sample have significant exposure. Griffin and Stulz (2001) use industry returns from developed markets and fail to find evidence of exposure. Several potential explanations exist for these weak results. One of the explanations pointed out by Bartov and Bodnar (1994) is that existing studies investigate almost exclusively linear/symmetric currency exposure and may fail to account for possible nonlinear/asymmetric relationship between the value of a firm and exchange rate, which motivates the current paper. In addition, theoretical literature suggests that firm behavior may well be different in periods of depreciation and appreciation, which should have an impact on how exchange rates affect the firm value.\(^2\) Although theoretically sounded, this asymmetric

\(^2\) Section II provides a brief review on the theoretical literature of asymmetric currency exposure.
response of stock returns to currency appreciations and depreciations has received very little
attention in the literature. Consequently, the current paper attempts to fill this gap by testing
asymmetric currency exposure empirically along with its pricing.

As far as the currency risk pricing is concerned, many studies have shown that
currency risk is priced in an international setting (see, e.g., Ferson et. al. (1987), Harvey
Gerard (1997, 1998)). However, whether it is priced in a domestic context is less conclusive.
For example, utilizing Ross’s (1976) APT model, Jorion (1991) finds that currency risk is not
priced in the US stock market during the period of 1971-87. Both Hamao (1988) and Brown
and Otsuki (1990) examine the pricing of currency risk in the Japanese market, and they also
conclude that it is not priced. These studies are all based on unconditional APT to draw their
conclusions, which may not be reliable since time-varying risk premium has been documented
in asset pricing literature. Ignoring this time-varying feature may bias the findings.
Consequently, using eight industry portfolios from Mexico, Bailey and Chung (1995) test a
conditional multi-factor model with time-varying ex ante risk premia. Their estimation results
show that none of the eight industry portfolios has significant currency exposure and that the
currency risk is not priced both unconditionally and conditionally. On the other hand, Choi et

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3 A recent paper by Koutmos and Martin (2003) has attempted to model asymmetric currency
exposure, but they only find 11.11% (7 out of 63) of their total sample with significant
asymmetric currency exposure and if considering the US data only, the percentage drops to
8.33% (3 out of 36). As a result, it demands another look at the asymmetric currency
exposure in addition to it pricing, which is ignored in Koutmos and Martin (2003).
al. (1998) test a conditional intertemporal CAPM allowing the prices of risk to change over time, and conclude that both bilateral and multilateral exchange rates are priced in the Japanese stock market, but their ‘pricing kernel’ approach does not allow them to estimate the risk exposure coefficients. As a result, whether (asymmetric) currency exposure is significant and whether it commands a risk premium in the domestic context such as US stock market are still unclear.

The current paper extends the existing literature in several important ways that provides new insights about the currency exposure and its pricing. First, I go beyond the traditional regression framework that imposes a linear relation between exchange rates and stock returns and allow for potentially different impacts of exchange rates during periods of depreciating versus appreciating currency values. That is, I investigate the possibility of asymmetric currency exposure that has not been fully explored in the literature. Second, I apply multivariate GARCH-in-mean (MGARCH-M) methodology to estimate asymmetric currency exposure coefficients. Most studies dealing with currency exposure use OLS or seeming unrelated regression (SUR). Without taking into account second moment temporal dependencies in asset returns, which have been documented extensively in literature (see, e.g., Hsieh (1989), Bollerslev et al. (1992), among others), both OLS and SUR will produce inefficient parameter estimates as well as biased test statistics, which may explain why previous studies have had difficulty finding significant currency exposure. The multivariate approach employed in this paper allows me to utilize the information in the entire variance-

4 The “pricing kernel” approach used by Choi et al. (1998) only allow them to model the price of factor risk.
covariance matrix of the errors, which, in turn, leads to more precise estimates of the parameters of the model. Also, many issues in finance can only be fully addressed within a multivariate framework. Third, I not only examine the presence of asymmetric currency exposure using MGARCH-M, but also test whether it is indeed priced in the stock market. In doing so, a more informative conclusion can be drawn with regard to the validity of corporate currency hedging discussed in the beginning. The utilization of MGARCH-M methodology overcomes the problems of two-step procedure usually employed by previous researchers when estimating factor models (see, e.g., Engle et al. (1990), Ng et al. (1992), Turtle et al. (1994), and Flannery et al. (1997)), and avoids the factor orthogonalization problem faced by previous researchers (see, e.g., Jorion (1991), Bailey and Chung (1995) and Choi et al. (1998)).5 Finally, to provide the robustness of the results, several additional tests are considered. Allayannis (1996) and Chow et al. (1997) argue that currency exposures increase as one extends the measurement interval. As a result, both weekly and monthly data are used to estimate the model. Because the main empirical results are derived using industry index returns, one may question the problem of industry level aggregation. The problem is that firms within an industry need not be homogeneous. It may be that industry-wide exposure is

5 In conformity with the original assumptions behind Ross’s (1976) APT, previous studies first use regressions to generate mean zero orthogonalized factors, and then use these generated factors to test factor models. As pointed out by Shanken (1992), this type of two-stage asset pricing test has an error-in-variable problem that can make the second-stage’s reported standard errors unreliable. The utilization of MGARCH-M in the current paper overcomes this problem because those orthogonalized factors will be created internally by the construction of MGARCH-M.
actually high but that individual firms within the industry are exposed in opposite ways. An aggregation of their returns will therefore average out the individual exposure effects (Dominguez and Tesar (2001)). Consequently, I re-estimate the model using securities from a single industry.

The empirical results show that 58% of US industries and 90% of US banks are asymmetrically affected by currency movements. In addition, the currency risk is priced regardless of whether bilateral or multilateral exchange rate is used, whether returns are measured at weekly or monthly horizon, and whether aggregate industry portfolios or individual securities are used to estimate the model. The strong empirical evidence found in this study implies that corporate currency hedging not only results in more stable cash flows for a firm, but also reduces its cost of capital, and hence is justifiable.\textsuperscript{6}

The remainder of the paper is organized as follows. Section II motivates the theoretical multi-factor asset pricing model. Section III presents the econometric methodologies used to test the model. Section IV discusses the data. Section V reports and discusses the empirical results. Concluding comments are offered in Section VI.

II. The Theoretical Motivation

\textsuperscript{6} Jorion (1991) finds that currency risk is not priced, and thus questions the justification of corporate currency hedging from asset-pricing viewpoint.
A. Theoretical Background on Asymmetric Currency Exposure

Several theoretical literature have been proposed to explain why firm behavior may well be different in periods of depreciation and appreciation, which should have an impact on how exchange rates affect the firm value. The theoretical explanations, which are summarized in Miller and Reuer (1998) and Koutmos and Martin (2003), include pricing-to-market (PTM), hysteretic behavior, and option theory.

A.1. Pricing-to-market

The asymmetric responses of stock prices to currency movements can be attributed to firms’ pricing-to-market behavior (Mann (1986), Giovannini (1988), Froot and Klemperer (1989), Knetter (1989, 1994), and Marston (1990)). Pricing-to-market essentially involves adjusting export prices based on the degree of competition in foreign exchange markets. Knetter (1994) considers two alternative scenarios that could give rise to asymmetric markups on exporter goods and therefore the asymmetric currency exposure. One scenario is limited distribution capacity or quotas in export markets, which could present a binding constraint on sales volume. Due to a sales volume constraint (SVC), exporters choose larger home-currency markups during periods of foreign currency appreciation than markdowns during periods of foreign currency depreciation. The second scenario is that exporters with market share objectives (MSO) may reduce home-currency margins during periods of foreign currency depreciation and maintain their profit margins by lowering the local currency price of the product during periods of foreign currency appreciation. As a result, the cash flow would increase to a lesser degree with foreign currency appreciations than decrease with its
deprecations. These two explanations for asymmetry between price markups and markdowns give rise to different patterns of exposure coefficients. In the case of SVC, home-currency margins will be constant during periods of foreign currency depreciation and increase during periods of foreign currency appreciation, and therefore we should observe no significant economic exposure to decreases in the dollar price of a foreign currency (US$/FX ↓ => no exposure), and a positive exposure to foreign currency appreciation (US$/FX ↑ => positive exposure). However, in the case of MSO, exporters maintain constant margins during periods of foreign currency appreciation and decrease their margins during periods of foreign currency depreciation. Consequently, we would observe no significant exposure to appreciation in the dollar value of a foreign currency (US$/FX ↑ => no exposure), and a positive exposure during periods of depreciation of the foreign currency (US$/FX ↓ => positive exposure).

### A.2. Hysteretic behavior

The asymmetric responses may also result from hysteretic behavior (Baldwin (1988), Baldwin and Krugman (1989), Dixit (1989), and Christophe (1997)). In this case, new export competitors are enticed to enter the market when the domestic currency depreciates, but their behavior is considered hysteretic if they still remain in the market once the currency appreciates. The reason why they continue to stay in the market is that the cost of reducing capital stock is higher than increasing it, which is the so-called irreversibility of investment (Baba and Fukao (2000)). As the new exporters enter the market when domestic currency depreciates, the cash flows of existing exporters may not increase to the degree that would
occur without the new entrants. However, if the new exporter continue to stay in the market as domestic currency strengthens, the cash flows of both ‘new’ and ‘existing’ exporters are likely to decrease. As a result, we would observe no exposure to appreciation in the dollar value of a foreign currency (US$/FX ↑ => no exposure), and a positive exposure during periods of depreciation of the foreign currency (US$/FX ↓ => positive exposure).

A.3. Asymmetric Hedging

Finally, the asymmetric responses may be due to asymmetric hedging behavior. Asymmetric hedging occurs when firms take one-sided hedges, such as with the use of currency options. Firms with net long positions may be willing to hedge against domestic currency appreciations yet remain unhedged against domestic currency depreciations. As a result, we would observe a positive exposure when foreign currency appreciates (US$/FX ↑ => positive exposure), and no exposure when foreign currency depreciates (US$/FX ↓ => no exposure). On the other hand, firms with net short positions are likely to hedge against domestic currency depreciations (US$/FX ↑ => no exposure), but remain unhedged against domestic currency appreciations (US$/FX ↓ => negative exposure). Thus, asymmetric hedging behavior produces an asymmetric impact on firms’ cash flows.

B. Multi-factor model with asymmetric currency exposure

We know that the first-order condition of any consumer-investor’s optimization problem can be written as:
\[ E[M_t R_{i,t} | \Omega_{t-1}] = 1, \ \forall i = 1 \cdots N \]  

(1)

where \( M_t \) is known as a stochastic discount factor or an intertemporal marginal rate of substitution; \( R_{i,t} \) is the gross return of asset \( i \) at time \( t \) and \( \Omega_{t-1} \) is market information known at time \( t-1 \). Without specifying the form of \( M_t \), equation (1) has little empirical content since it is easy to find some random variable \( M_t \) for which the equation holds. Thus, it is the specific form of \( M_t \) implied by an asset pricing model that gives equation (1) further empirical content (see, e.g., Ferson (1995)). Suppose \( M_t \) and \( R_{i,t} \) have the following factor representations:

\[
M_t = \alpha_0 + \sum_{k=1}^{K} \beta_{ik} F_{k,t} + u_t
\]  

(2)

\[
r_{i,t} = \alpha_i + \sum_{k=1}^{K} \beta_{ik} F_{k,t} + \varepsilon_{i,t} \ \forall i = 1 \cdots N
\]  

(3)

where \( r_{i,t} = R_{i,t} - R_{0,t} \) is the raw returns of asset \( i \) in excess of the risk-free rate, \( R_{0,t} \), at time \( t \), and \( E[u_t F_{k,t} | \Omega_{t-1}] = E[u_t | \Omega_{t-1}] = E[\varepsilon_{i,t} F_{k,t} | \Omega_{t-1}] = E[\varepsilon_{i,t} | \Omega_{t-1}] = E[\varepsilon_{i,t} u_t | \Omega_{t-1}] = 0 \ \forall i, k \); \( F_{k,t} \) is a common risk factor which captures systematic risk affecting all assets \( r_{i,t} \) including \( M_t \); \( \beta_{ik} \) is the associated time-invariant factor exposure which measures the sensitivity of the asset \( i \) to the common risk factor \( k \), while \( u_t \) is an innovation and \( \varepsilon_{i,t} \) is an idiosyncratic
term which reflects unsystematic risk. The risk-free rate, \( R_{0,t-1} \), must also satisfy equation (1).

\[
E[M_t R_{0,t-1} | \Omega_{t-1}] = 1
\]  

(4)

Subtract equation (4) from equation (1), obtaining:

\[
E[M_t r_{i,t} | \Omega_{t-1}] = 0 \quad \forall i = 1 \ldots N
\]  

(5)

Apply the definition of covariance to equation (5), obtaining:

\[
E[r_{i,t} | \Omega_{t-1}] = \frac{Cov(r_{i,t} ; -M_j | \Omega_{t-1})}{E[M_j | \Omega_{t-1}]} \quad \forall i = 1 \ldots N
\]  

(6)

Substituting the factor model in equation (3) into the right hand side of equation (6) and assuming that \( Cov(\varepsilon_{i,t}; M_{i+1}) = 0 \) implies

\[
E[r_{i,t} | \Omega_{t-1}] = \sum_k \beta_{ik} \left[ \frac{Cov(F_{k,t}; -M_j | \Omega_{t-1})}{E[M_j | \Omega_{t-1}]} \right] = \sum_k \beta_{ik} \lambda_{k,t-1} \quad \forall i = 1 \ldots N
\]  

(7)

\[7\] The empirical studies of Ferson and Harvey (1993) and Ferson and Korajczyk (1995) consistently show that movements in factor exposures/betas account for only a small fraction of the predictable change in expected returns, in both the domestic and the international context. Thus, to simplify the model, a time-invariant factor beta seems to be reasonable.
where $\lambda_{k,t-1}$ is the time-varying risk premium per unit of beta risk.

Assuming $F_{k,t} = E[F_{k,t} \mid \Omega_{t-1}] + \varepsilon_{k,t}$, where $\varepsilon_{k,t}$ is the factor innovation with $E[\varepsilon_{k,t} \mid \Omega_{t-1}] = 0$, then equation (3) can be rewritten as:

$$r_{i,t} = \alpha_i + \sum_k \beta_{ik} E[F_{k,t} \mid \Omega_{t-1}] + \sum_k \beta_{ik} \varepsilon_{k,t} + \varepsilon_{i,t} \quad \forall i = 1, \ldots, N$$

(8)

Taking conditional expectation on both sides of equation (8) and compare it with equation (7), then under the null hypothesis of $\alpha_i = 0$ obtaining:

$$E[F_{k,t} \mid \Omega_{t-1}] = \lambda_{k,t-1} \quad \forall k$$

(9)

Substituting equation (9) into equation (8), and assuming $\alpha_i = 0$, equation (8) becomes

$$r_{i,t} = \sum_k \beta_{ik} (\lambda_{k,t-1} + \varepsilon_{k,t}) + \varepsilon_{i,t} \quad \forall i = 1, \ldots, N$$

(10)

Since the main focus of the current paper is to test the existence of asymmetric currency exposure and its price, I consider a two-factor model (i.e., $k = 2$ in Equation (10)) where the two factors are world market and currency risks. To incorporate the asymmetry of currency exposure in the model, I can rewrite equation (10) as:

$$r_{i,t} = (\lambda_{m,t} + \varepsilon_{m,t}) \beta_{im} + (\lambda_{c,t} + \varepsilon_{c,t})(\beta_{ic} + \beta_{ic}^D D_{i,t}) + \varepsilon_{i,t} \quad \forall i = 1, \ldots, N$$

(11)
where “$m$” denotes world market risk, and “$c$” denotes currency risk. $D_t$ is a dummy variable, which is equal to one if $\varepsilon_{c,t} < 0$ and zero otherwise.\footnote{Previous studies (e.g., Jorion (1990, 1991), Bodnar and Gentry (1993), Choi and Prasad (1995), Chow et al. (1997), He and Ng (1998), Koutmos and Martin (2003), among others) on currency exposure usually assume that the first differences of exchange rate are unanticipated, and thus can be treated as innovations to estimate currency exposures. However, in this paper the unanticipated components of exchange rate changes are the residuals obtained from the MGARCH-M model, so not only are these residuals truly unexpected by construction, but also they are uncorrelated with other factors based on the multivariate nature of GARCH processes, which avoids the error-in-variable problem when using two-step procedure to create orthogonalized factors.} For a given value of market portfolios, the response of $r_{i,t}$ will be equal to $\beta_{ic}$ when $\varepsilon_{c,t} > 0$ and $\beta_{ic} + \beta_{ic}^d$ for $\varepsilon_{c,t} < 0$.

Equation (11) can be used to test the null hypothesis that currency exposure is symmetric, i.e., $H_0 : \beta_{ic}^d = 0$. A measure of the degree of asymmetric response of excess returns to currency movements can be constructed by taking the ratio $\beta_{ic}^d / \beta_{ic}$. 

III. Econometric Methodology

A. Modeling Conditional Factor Risk Premia
Theoretical work by Merton (1973) relates the expected risk premium of factor \( k \), \( \lambda_{k,t-1} \), in equation (10) to its volatility and a constant proportionality factor. In supporting these theoretical results, Metron (1980) tests a single-beta market model and find that the expected risk premium on the stock market is positively correlated with the predictable volatility of stock returns.\(^9\) As a result, the following relationship is postulated for \( \lambda_{k,t-1} \) \( \forall k \) :

\[
\lambda_{k,t-1} = E(F_{k,t} \mid \Omega_{t-1}) = k_0 + k_1 h_{k,t} \quad \forall k
\]

(12)

where \( h_{k,t} \) is factor \( k \)'s conditional variance, and \( k_0 \) and \( k_1 \) are constant parameters. To complete the conditional two-factor model with time-varying risk premia, equation (11) can be rewritten as:

\[
r_{i,t} = (m_0 + m_1 h_{m,t} + \varepsilon_{m,t}) \beta_{i,m} + (c_0 + c_1 h_{e,t} + \varepsilon_{e,t}) (\beta_{i,c} + \beta_{i,c}^d D_t) + \varepsilon_{i,t} \quad \forall i
\]

(13)

Equation (13) can be used to test whether the predictable volatilities of the market-wide risk factors are significant sources of risk in addition to the test of the significance of the asymmetric currency exposure.

B. Modeling Conditional Variance-Covariance Structure

\(^9\) Alternatively the factor risk premium can be constructed by using ‘information’ variables similar to Ferson and Harvey (1991, 1993).
The conditional two-factor asset pricing model in equation (13) has to hold for every asset. However, the model does not impose any restrictions on the dynamics of the conditional second moments. Given the computational difficulties in estimating a larger system of asset returns, parsimony becomes an important factor in choosing different parameterizations. A popular parameterization of the dynamics of the conditional second moments is BEKK, proposed by Baba, Engle, Kraft, and Kroner (1989). The major feature of this parameterization is that it guarantees that the variance-covariance matrices in the system are positive definite. However, it still requires researchers to estimate a larger number of parameters. Instead of using BEKK specification, I employ a parsimonious parameterization of the conditional variance-covariance structure of asset returns and risk factors proposed by Ding and Engle (1994). Their parameterization allows me to reduce the number of parameters to be estimated significantly.\(^\text{10}\)

Under Ding and Engle’s parameterization, the conditional second moments is assumed to follow a diagonal process and the system is assumed to be covariance stationary; therefore, the GARCH process for the conditional variance-covariance matrix of asset returns and risk factors can be written as,

\[
H_t = H_0 * \left( (u^T - aa^T - bb^T) + a a^T \epsilon_{t-1}^T \epsilon_{t-1} + b b^T H_{t-1} \right)
\]

\[\text{(14)}\]

\(\text{10}\) In a diagonal system with \(N\) assets and \(K\) factors, the number of unknown parameters in the conditional variance equation is reduced from \(2(N + K)^2 + \frac{(N + K)(N + K + 1)}{2}\) under BEKK specification to \(2(N + K)\) under Ding and Engle’s specification.
where $H_t \in \mathbb{R}^{(N+K)\times(N+K)}$ is a time-varying variance-covariance matrix of asset returns and risk factors. $N+K$ is the number of equations where the first $N$ equations are those for the asset returns and the last $K$ equations are those for the risk factors. $H_0$ is the unconditional variance-covariance matrix of unsystematic risks and factor innovations. $1$ is a $(N+K)\times1$ vector of ones, $a,b \in \mathbb{R}^{(N+K)\times1}$ are vectors of unknown parameters, and $*$ denotes element-by-element matrix product. The $H_0$ is unobservable and has to be estimated. As suggested by De Santis and Gerard (1997, 1998), it can be consistently estimated using iterative procedure. In particular, $H_0$ is set equal to the sample covariance matrix of the asset returns and risk factors in the first iteration, and then it is updated using the covariance matrix of the estimated residual at the end of each iteration.

Under the assumption of conditional normality, the log-likelihood to be maximized under both processes can be written as,

$$\ln L(\theta) = -\frac{T\times(N+K)}{2} \ln 2\pi - \frac{1}{2} \sum_{t=1}^T \ln |H_t(\theta)| - \frac{1}{2} \sum_{t=1}^T \varepsilon_t(\theta)^T H_t(\theta)^{-1} \varepsilon_t(\theta)$$  \hspace{1cm} (15)$$

where $\theta$ is the vector of unknown parameters in the model. Since the normality assumption is often violated in financial time series, the quasi-maximum likelihood estimation (QML) proposed by Bollerslev and Wooldridge (1992) which allows inference in the presence of departures from conditional normality is employed. Under standard regularity conditions, the QML estimator is consistent and asymptotically normal and statistical inferences can be carried out by computing robust Wald statistics. The QML estimates can be obtained by
maximizing equation (15), and calculating a robust estimate of the covariance of the parameter estimates using the matrix of second derivatives and the average of the period-by-period outer products of the gradient. Optimization is performed using the Broyden, Fletcher, Goldfarb and Shanno (BFGS) algorithm, and the robust variance-covariance matrix of the estimated parameters is computed from the last BFGS iteration.

IV. Data and Summary Statistics

I analyze the currency exposure and its pricing at the industry level. Industry total return (including dividends) indices are collected from Datastream at a weekly frequency over the period August 4, 1978 to December 28, 2001. A major strength of this data source is that Datastream classifies industry indices into one of six levels, and at each additional level there are more disaggregated industry definitions until the most disaggregated industry classification, level 6. As pointed out by Griffin and Karolyi (1998), using broad industrial classifications leads to lumping together heterogeneous industries, and thus to increase the power of detecting priced factors, a more disaggregated industry data such as those obtained from Datastream is desirable. There are 72 level 6 industries for the US, and since this study applies MGARCH-M methodology to estimate the model, the number of assets chosen to estimate the model becomes a concern. As a result, to keep the dimension of variance-covariance matrix of asset returns manageable, twelve industry indices are examined in this study. They are airlines & airports (AIRLN), banks (BANKS), electricity (ELECT), general industrials (GENIN), hotels (HOTEL), media & photograph (MEDIA), automobiles
(AUTOS), chemicals (CHMCL), electronic equipment (ELETR), electrical equipment (ELTNC), paper products (PAPER), and pharmaceuticals (PHARM). Datastream world total return index (TOTMK) is used to construct the world market risk. Both bilateral and multilateral exchange rates are used to construct the currency risk factors ($c_F$) and to estimate asymmetric currency exposures. The bilateral exchange rate is the Japanese yen (JPY) against the US dollar (USD), expressed as USD/JPY.\textsuperscript{11} The multilateral rate is the trade-weighted average of the foreign exchange values of the USD against the currencies of a large group of major US trading partners (TWFX). In all instances, the currency is expressed as US dollar price per unit of foreign currency, so a positive change indicates a decreasing value of the USD. Finally, seven-day Eurodollar deposit rate is used to compute excess industry and market returns. All the data are extracted from Datastream. Table 1 describes the variables and notations used in this paper.

Table 2 presents some descriptive statistics of the continuously compounded weekly returns on the industry indices, world market index, and bilateral and multilateral exchange rates. As can be seen from Panel A, the ELTNC has the highest mean returns (0.191%), while AIRLN has the lowest mean return (0.019%) among all industry indices. For the exchange rates, the mean of log first differences of USD/JPY is positive (0.03%), suggesting that USD was decreasing in value against JPY over the sample period. For multilateral exchange rate, the average of log first differences of TWFX is negative (-0.011%), implying that the USD

\footnote{In empirical tests, I also consider other major bilateral exchange rates including British pound and European Currency Unit. The results not reported here are very similar, and are available upon request.}
was appreciating on average against its trading partners.

Table 2 also reports skewness and the index of excess kurtosis. The distribution of the returns in all instances show significant excess kurtosis suggesting that the return series are conditionally heteroskedastic. The use of GARCH model will be able to take that into account. The correlation coefficients between industry returns and world market returns are all positive, suggesting that all the industries in the sample are positively exposed to world market risk. On the other hand, most of the industries (9 out of 12) are negatively correlated with USD/JPY bilateral exchange rate, indicating that those industries have negative exposures to the depreciation of the USD.

V. Empirical Results

A. Exposure to a Bilateral exchange rate

Table 3 reports the estimation results for the conditional two-factor asset pricing model (equation (13)) where the currency risk is proxied by the USD/JPY bilateral exchange rate. First consider the risk exposure coefficients, as can be seen in the table, 50% (6 out of 12) of world market exposures are significant. For currency exposure, only 25% of industries (3 out of 12) have significant $\beta_c$, but 58% of the industries (7 out of 12) have significant $\beta_{c^d}$, indicating strong evidence of asymmetry in currency exposures for US industries with respect to USD/JPY exchange rate. This is an extremely important finding as it suggests that
models assume symmetric exposure over appreciation-depreciation cycles are frequently misspecified. Among the 7 industries with documented significantly negative $\beta_{c}^{d}$, 2 of them (BANKS and PHARM) have significantly positive $\beta_{c}$, suggesting that the type of asymmetry is consistent with asymmetric hedging behavior and/or PTM with SVC if they are net exporters. For the other five industries (GENIN, MEDIA, CHMCL, ELETR, and ELTNC), the asymmetry is compatible with asymmetric hedging behavior and/or PTM with SVC if they are net importers because all of them have insignificant $\beta_{c}$. The overall impact of currency movements on returns can be assessed by adding the estimated $\beta_{c}$ to $\beta_{c}^{d}$ coefficients. From Panel A of Table 7 the estimated total currency exposures are all negative except AIRLN, suggesting that for a given value of market portfolio US industries suffer from unexpected USD depreciations since the currency factor is measured in US dollar per unit of foreign currency. To measure the degree of asymmetric response of industry returns to USD/JPY movements, I also calculate the ratio of $\beta_{c}^{d} / \beta_{c}$ for all industries. As can be seen from Panel A of Table 7, all the ratios are greater than one in absolute value except PAPER, implying strong asymmetric exposure with respect to USD/JPY movements.

As far as the time-varying factor risk premium is concerned, strong evidence of GARCH in mean effects is found for the dynamics of both world market and currency risks since the parameters $(m_{1}, c_{1})$ are all significant at the 1% level, and thus they have significant impact on the dynamics of factor risk premia. This finding implies that not only is USD/JPY exposure significant, but it is priced in the US stock market, which sheds a new light on the issue of currency risk pricing in a domestic context since Jorion (1991) concludes that
currency risk is not priced in the US stock market using industry indices. The strong evidence of factor sensitivities and time-varying world market and currency risk premia found in this paper points out the advantage of using MGARCH-M approach over the traditional OLS or SUR approaches where nonlinear second moment dependencies are ignored.

Next, consider the estimated parameters for the conditional variance-covariance processes. All of the elements in the vectors $a$ and $b$ are statistically significant at the 1% level, implying the strong presence of GARCH effect in all return series. In addition, the estimates satisfy the stationarity conditions for all the variance and covariance processes.\(^{12}\) The presence of conditional heteroskedasticity and the high degree of volatility persistence suggest that using simple OLS or SUR, which assume constant variance, will lead to high standard errors and erroneous inferences.\(^{13}\) This may help to explain why previous studies have failed to detect both significant currency exposure and currency risk price.

### B. Exposure to a Multilateral Currency Index

To check the robustness of the result of significant asymmetric exposures using bilateral rate, in the section I use trade-weighted exchange rate (TWFX) to measure

\(^{12}\) For the process in $H$ to be covariance stationary, the condition $a_i a_j + b_i b_j < 1 \ \forall i, j$ has to be satisfied. (see, e.g., Bollerslev (1986), and De Santis and Gerard (1997, 1998))

\(^{13}\) The degree of volatility persistence can be obtained by comparing the parameter estimates $a$ and $b$ in the GARCH process. Since most of the $b$ estimates are in the range of 0.96 to 0.97, the volatility is highly persistence.
asymmetric exposure. Table 4 reports the estimation results. As can be seen, only 3 out of the 12 industries (BANKS, MEDIA, and PHARM) are significantly exposed to TWFX risk, and two of them (BANKS and PHARM) along with ELECT are exposed asymmetrically, suggesting that the both symmetric and asymmetric currency exposures with respect to TWFX are not very strong for US industries. This weak result is consistent with previous studies using trade-weighted exchange rate index (e.g., Jorion (1990, 1991)), and the point made by Dominguez and Tesar (2001) that the use of the trade-weighted exchange rate index is likely to understate the extent of exposure. In terms of the sign of exposures, the source of asymmetry is compatible with asymmetric hedging. Alternatively the asymmetry could be the result of asymmetric PTM behavior with sales volume constraint. Regarding the total currency exposure, the sum of $\beta_c + \beta_c^d$ are all negative (see Panel B of Table 7) for the 3 industries with documented significant asymmetric exposures (BANKS, ELECT, and PHARM), indicating that the US industries also suffer from unexpected TWFX depreciations. The ratio of $\beta_c^d / \beta_c$ is greater than one in absolute value, so the degree of asymmetry is very high for each of the 3 industries.

For factor risk premium, both world market and currency risks are again not only priced, but change over time since the parameters $(m_1, c_1)$ are statistically significant at the 1% level. The strong evidence of time-varying currency risk premium using the trade-weighted exchange rate index suggests that the failure of previous studies (e.g., Jorion (1991)) in testing currency risk pricing in the US stock market may not be necessarily due to which exchange rate is used, but possibly due to the ignorance of conditional heteroskedasticity and the joint estimation of currency exposure and the price of currency risk.
C. Sensitivity of Exposure to Horizon

Several studies of exposure have found that the extent of estimated exposure is increasing in the return horizon (see, for example, Bartov and Bodnar (1994), Alayannis (1996), Bodnar and Wong (2000) and Chow et al. (1997)). Indeed, most studies of exposure are conducted using monthly returns, suggesting that the results based on weekly returns presented earlier may understate the true extent of exposure. Therefore, in this section I re-estimate the model using monthly returns for the same 12 industries and bilateral USD/JPY exchange rate. Table 5 shows the estimation results. The world market exposure, $\beta_m$, is significant in 75% of the cases (9 out of 12). For currency exposure, it is significant in 3 cases for $\beta_c$ and four cases for $\beta^d_c$. Compare with the results using weekly returns, the number of significant asymmetric currency exposure has dropped from 7 cases (58%) to 4 cases (33%). This finding is in contrast to the literature that exposure is increasing in the return horizon (Bartov and Bodnar (1994), Alayannis (1996), Bodnar and Wong (2000) and Chow et al. (1997)), but may not be so surprising since MGARCH-M instead of OLS/SUR is used to estimate exposures and it is well known that GARCH effect diminishes as the return interval increases. Regarding the total currency exposures and the degree of asymmetry, Panel C of Table 7 shows that the results are virtually unchanged compared with the results found in Section A when using weekly data. That is, US industries benefit from a strong domestic currency, and the degree of asymmetry is very strong for those industries with documented significant asymmetric exposures. Finally, consistent with previous sections both world market and currency risks are significantly priced at the 1% level.
D. Exposure at the Individual Firm Level

As mentioned in the Introduction that the problem of industry level aggregation is that firms within an industry need not be homogeneous. It may be that industry-wide exposure is actually high but that individual firms within the industry are exposed in opposite ways. An aggregation of their returns will therefore average out the individual exposure effects (Dominguez and Tesar (2001)). Consequently, I re-estimate the model using securities from a single industry - bank.\footnote{While exchange rate can influence the value of firms in many industries, my focus on banks stems from the growing international interest in monitoring banks’ market risk, including currency risk after the occurrences of several financial crises in late 1990s.} Total stock return (dividend included) indices for ten US major banks obtained from Datastream are selected to estimate the model. The ten banks are Bank of America (BAC), Wells Fargo (WFC), Bank of New York (BK), Mellon Financial (MEL), National City (NCC), The PNC Financial Services (PNC), Wachovia (WB), Comerica (CMA), Northern Trust (NTRS), Riggs National (RIGS). The estimation results reported in Table 6 show that 70% of the bank stocks (7 out of 10) are significantly exposed to USD/JPY movements, and almost all the banks (9 out of 10) are exposed asymmetrically. This strong evidence of symmetric and asymmetric currency exposures implies that exposures are much stronger at the firm level than at the industry level. Since $\beta^d_c$ is negative for all industries and $\beta_c$ is positive for all industries except RIGS, the type of asymmetry is consistent with asymmetric hedging behavior. To the extent that US banks asymmetrically hedge their net exposure to foreign currencies, this finding makes sense since in addition to being the major suppliers of hedging instruments to their clients, US banks are also sophisticated users on

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their own accounts. We can also calculated the total currency exposure \((\beta_c^d + \beta_c^e)\) and the ratio of \(\beta_c^d / \beta_c^e\) for each bank. It is not difficult to find that \(\beta_c^d + \beta_c^e\) is negative for all the banks, suggesting that US banks also suffer from the unexpected US dollar depreciations, and \(\beta_c^d / \beta_c^e\) are all greater than one in absolute value except RIGS, indicating strong asymmetry at the firm level.

Regarding the factor risk premium, both market and currency risk premia are once again significant at the 1% level, indicating that USD/JPY currency risk is priced not only in the aggregate portfolios, but also in individual securities, which sheds a new light on the pricing of currency risk in US stock market.\(^{15}\) The strong evidence of asymmetric currency exposure and its pricing found in this section confirms previous results using aggregate industry portfolios.

VI. Conclusion

Whether stock returns are linked to currency movements and whether currency risk is priced in a domestic context are two inter-related issues, which are critical to investors and corporate risk management. Surprisingly many previous studies have examined these two issues in isolation, and the results are less conclusive and thus still subject to a great debate. In this paper, I attempt to provide new empirical evidence on these two inter-related issues

\(^{15}\) Using US industry portfolios, Jorion (1991) is not able to detect significant currency risk premium.
using a different approach. In particular, I utilize the multivariate GARCH in mean approach to estimate a conditional two-factor asset pricing model. The advantage of this approach is that it allows me to simultaneously estimate both the risk pricing parameters and risk exposure coefficients within a conditional multivariate framework.

The empirical results show that 58% of US industries and 90% of US banks are asymmetrically affected by USD/JPY movements. In addition, both US industries and banks benefit from a strong domestic currency. Finally, the currency risk is priced regardless of whether bilateral or multilateral exchange rate is used, whether returns are measured at weekly or monthly horizon, and whether aggregate industry portfolios or individual securities are used to estimate the model.16 The strong evidence found here sheds a new light on the issues of currency exposure and the pricing of currency risk in the US stock market since previous studies are not successful in detecting them. Two reasons may explain why they fail. First, the conditional heteroskedasticity is not explicitly modeled. Second, asymmetric currency exposure is ignored.

16 Although in this paper I mainly focus on the US market since Jorion (1990, 1991) conclude that currency risk exposure is not significant and is not priced in the US stock market, the MGARCH-M approach adopted here can be easily applied to test asymmetric exposure for other countries.
References


**TABLE 1**
Variable Definitions and Notations

<table>
<thead>
<tr>
<th>Industry</th>
<th>Risk factor</th>
</tr>
</thead>
<tbody>
<tr>
<td>Airlines &amp; Airports</td>
<td>Datastream world total return index</td>
</tr>
<tr>
<td>Banks</td>
<td>USD/JPY</td>
</tr>
<tr>
<td>Electricity</td>
<td>Trade-weighted US dollar exchange rate</td>
</tr>
<tr>
<td>General Industrials</td>
<td>TOTMK</td>
</tr>
<tr>
<td>Hotels</td>
<td>USD/JPY</td>
</tr>
<tr>
<td>Media &amp; Photograph</td>
<td>TWFX</td>
</tr>
<tr>
<td>Automobiles</td>
<td>USD/JPY</td>
</tr>
<tr>
<td>Chemicals</td>
<td>TWFX</td>
</tr>
<tr>
<td>Electronic Equipment</td>
<td>USD/JPY</td>
</tr>
<tr>
<td>Electrical Equipment</td>
<td>TWFX</td>
</tr>
<tr>
<td>Paper</td>
<td>USD/JPY</td>
</tr>
<tr>
<td>Pharmaceuticals</td>
<td>USD/JPY</td>
</tr>
<tr>
<td></td>
<td>USD/JPY</td>
</tr>
</tbody>
</table>

*Industries and Risk factors are represented by their respective notations.*
TABLE 2
Summary Statistics and Correlation Coefficients

<table>
<thead>
<tr>
<th></th>
<th>AIRLN</th>
<th>BANKS</th>
<th>ELECT</th>
<th>GENIN</th>
<th>HOTEL</th>
<th>MEDIA</th>
<th>AUTOS</th>
<th>CHMCL</th>
<th>ELETR</th>
<th>ELTNC</th>
<th>PAPER</th>
<th>PHARM</th>
<th>USD/JPY</th>
<th>TWFX</th>
<th>TOTMK</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean (%)</td>
<td>0.019</td>
<td>0.170</td>
<td>0.089</td>
<td>0.134</td>
<td>0.111</td>
<td>0.103</td>
<td>0.061</td>
<td>0.103</td>
<td>0.157</td>
<td>0.191</td>
<td>0.033</td>
<td>0.204</td>
<td>0.030</td>
<td>-0.011</td>
<td>0.085</td>
</tr>
<tr>
<td>Skewness</td>
<td>-0.749**</td>
<td>0.020</td>
<td>-0.296**</td>
<td>-0.609**</td>
<td>-0.891**</td>
<td>-0.436**</td>
<td>-0.117</td>
<td>-0.359**</td>
<td>-0.243**</td>
<td>-0.230**</td>
<td>-0.370**</td>
<td>-0.401**</td>
<td>1.019**</td>
<td>-0.229**</td>
<td>-0.545**</td>
</tr>
<tr>
<td>Kurtosis</td>
<td>8.238**</td>
<td>3.394**</td>
<td>4.544**</td>
<td>5.256**</td>
<td>7.440**</td>
<td>2.767**</td>
<td>1.925**</td>
<td>3.586**</td>
<td>7.786**</td>
<td>3.010**</td>
<td>4.788**</td>
<td>1.497**</td>
<td>7.190**</td>
<td>1.453**</td>
<td>4.374**</td>
</tr>
</tbody>
</table>

The top half of this table presents summary statistics for the weekly dollar-denominated excess industry and world market returns and the log first differences of bilateral and trade-weighted exchange rates from 08/11/1978 through 12/28/01. The pairwise correlation coefficients between equity returns and currency changes are depicted in the bottom half.
TABLE 3

<table>
<thead>
<tr>
<th>Factor risk exposure</th>
<th>Conditional Mean Process</th>
<th>Conditional Variance Process</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\beta_m$</td>
<td>$\beta_c$</td>
</tr>
<tr>
<td>AIRLN</td>
<td>0.058 (0.046)</td>
<td>0.011 (0.059)</td>
</tr>
<tr>
<td>BANKS</td>
<td>-0.004 (0.033)</td>
<td>0.129 (0.041)**</td>
</tr>
<tr>
<td>ELECT</td>
<td>0.029 (0.025)</td>
<td>0.047 (0.037)</td>
</tr>
<tr>
<td>GENIN</td>
<td>0.045 (0.012)**</td>
<td>0.023 (0.015)</td>
</tr>
<tr>
<td>HOTEL</td>
<td>0.036 (0.060)</td>
<td>0.150 (0.075)**</td>
</tr>
<tr>
<td>MEDIA</td>
<td>0.010 (0.031)</td>
<td>0.067 (0.035)</td>
</tr>
<tr>
<td>AUTOS</td>
<td>0.042 (0.047)</td>
<td>0.031 (0.053)</td>
</tr>
<tr>
<td>CHMCL</td>
<td>-0.042 (0.021)*</td>
<td>0.013 (0.017)</td>
</tr>
<tr>
<td>ELETR</td>
<td>0.098 (0.045)*</td>
<td>0.091 (0.068)</td>
</tr>
<tr>
<td>ELTNC</td>
<td>0.077 (0.012)**</td>
<td>0.015 (0.021)</td>
</tr>
<tr>
<td>PAPER</td>
<td>0.081 (0.028)**</td>
<td>-0.025 (0.064)</td>
</tr>
<tr>
<td>PHARM</td>
<td>-0.215 (0.057)**</td>
<td>0.119 (0.021)**</td>
</tr>
</tbody>
</table>

Estimations are based on weekly dollar-denominated excess returns from 08/11/1978 through 12/28/01. Each mean equation relates the excess industry returns $r_{it}$ to its world market and USD/JPY currency risks. The factor realizations, $F_{m,t}$ and $F_{c,t}$, depend on its own conditional volatility, and thus the expected factor risk premia, $\lambda_{m,t}$ and $\lambda_{c,t}$, are the functions of conditional volatility.

$$r_{it} = (\lambda_{m,t} + \epsilon_{m,t}) \beta_{m} + (\lambda_{c,t} + \epsilon_{c,t})(\beta_{c} + \beta_{d}^{c} D_{t}) + \epsilon_{t} \quad \forall i$$

where $F_{m,t} = E(F_{m,t} | \Omega_{t-1}) + \epsilon_{m,t} = \lambda_{m,t} + \epsilon_{m,t} = m_0 + m h_{m,t} + \epsilon_{m,t}$

$$F_{c,t} = E(F_{c,t} | \Omega_{t-1}) + \epsilon_{c,t} = \lambda_{c,t} + \epsilon_{c,t} = c_0 + c h_{c,t} + \epsilon_{c,t}$$

$D_t$ is a dummy variable, which is equal to one if $\epsilon_{c,t} < 0$ and zero otherwise.

$$\epsilon_{t} | \Omega_{t-1} \sim N(0, H_{t})$$

The conditional covariance matrix $H_{t}$ is parameterized as follows

$$H_{t} = H_0 * (t^T - aa^T - bb^T) + aa^T * \epsilon_{t-1} \epsilon_{t-1}^T + bb^T * H_{t-1}$$

where $H_{t} \in \mathbb{R}^{14 \times 14}$ is the conditional covariance matrix of twelve industry returns and two risk factors. The elements of vectors $a, b \in \mathbb{R}^{14 \times 1}$ are the GARCH parameters, $t$ is a 14 x 1 unit vector and * denotes the Hadamard product (element-by-element multiplication). QML standard errors are reported in parentheses. * and ** denote statistical significance at the 5% and 1% level, respectively.
TABLE 4  

<table>
<thead>
<tr>
<th>Factor risk exposure</th>
<th>Conditional Mean Process</th>
<th>Conditional Variance Process</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\beta_m$</td>
<td>$\beta_c$</td>
</tr>
<tr>
<td>AIRLN</td>
<td>0.073 (0.041)</td>
<td>-0.117 (0.110)</td>
</tr>
<tr>
<td>BANKS</td>
<td>0.001 (0.033)</td>
<td>0.190 (0.065)**</td>
</tr>
<tr>
<td>ELECT</td>
<td>0.045 (0.026)</td>
<td>0.104 (0.062)</td>
</tr>
<tr>
<td>GENIN</td>
<td>0.030 (0.010)**</td>
<td>0.015 (0.023)</td>
</tr>
<tr>
<td>HOTEL</td>
<td>0.035 (0.054)</td>
<td>0.178 (0.118)</td>
</tr>
<tr>
<td>MEDIA</td>
<td>0.005 (0.025)</td>
<td>0.102 (0.047)*</td>
</tr>
<tr>
<td>AUTOS</td>
<td>0.026 (0.043)</td>
<td>-0.014 (0.081)</td>
</tr>
<tr>
<td>CHMCL</td>
<td>-0.041 (0.013)**</td>
<td>0.000 (0.026)</td>
</tr>
<tr>
<td>ELETR</td>
<td>0.058 (0.064)</td>
<td>0.119 (0.090)</td>
</tr>
<tr>
<td>ELTNC</td>
<td>0.061 (0.016)**</td>
<td>0.012 (0.028)</td>
</tr>
<tr>
<td>PAPER</td>
<td>0.086 (0.015)**</td>
<td>-0.141 (0.088)</td>
</tr>
<tr>
<td>PHARM</td>
<td>-0.106 (0.031)**</td>
<td>0.136 (0.069)*</td>
</tr>
</tbody>
</table>

Factor risk premium

| TOTMK | $m_0$ | 7.32E-05 (0.000) | 0.188 (0.005)** | 0.960 (0.003)** |
|       | $m_1$ | 0.015 (0.004)** | 0.156 (0.015)** | 0.948 (0.018)** |
| TWFX  | $c_0$ | 0.000 (0.000) | 0.156 (0.015)** | 0.948 (0.018)** |
|       | $c_1$ | -0.009 (0.003)** | 0.156 (0.015)** | 0.948 (0.018)** |

Estimations are based on weekly dollar-denominated excess returns from 08/11/1978 through 12/28/01. Each mean equation relates the excess industry returns $r_{it}$ to its world market and TWFX currency risks. The factor realizations, $F_{m,t}$ and $F_{c,t}$, depend on its own conditional volatility, and thus the expected factor risk premia, $\lambda_{m,t}$ and $\lambda_{c,t}$, are the functions of conditional volatility.

$r_{it} = (\lambda_{m,t} + \epsilon_{m,t})\beta_{m} + (\lambda_{c,t} + \epsilon_{c,t})(\beta_{c} + \beta_{d}^cD_t) + \epsilon_{t,i}$ \hspace{1cm} \forall i$

where $F_{m,t} = E(F_{m,t} | \Omega_{t-1}) + \epsilon_{m,t} = \lambda_{m,t} + \epsilon_{m,t} = m_0 + m_1h_{m,t} + \epsilon_{m,t}$

$F_{c,t} = E(F_{c,t} | \Omega_{t-1}) + \epsilon_{c,t} = \lambda_{c,t} + \epsilon_{c,t} = c_0 + c_1h_{c,t} + \epsilon_{c,t}$

$D_t$ is a dummy variable, which is equal to one if $\epsilon_{c,t} < 0$ and zero otherwise.

$\epsilon_t | \Omega_{t-1} \sim N(0, H_t)$

The conditional covariance matrix $H_t$ is parameterized as follows

$H_t = H_0 \ast (\mathbf{u}^T - \mathbf{a}^T \ast \mathbf{bb}^T) + \mathbf{aa}^T \ast \mathbf{e}_{t-1}e_{t-1}^T + \mathbf{bb}^T \ast H_{t-1}$

where $H_t \in \mathbb{R}^{14 \times 14}$ is the conditional covariance matrix of twelve industry returns and two risk factors. The elements of vectors $a, b \in \mathbb{R}^{14 \times 1}$ are the GARCH parameters, $\mathbf{t}$ is a 14 x 1 unit vector and $\ast$ denotes the Hadamard product (element-by-element multiplication). QML standard errors are reported in parentheses. * and ** denote statistical significance at the 5% and 1% level, respectively.
**TABLE 5**

<table>
<thead>
<tr>
<th>Factor risk exposure</th>
<th>Conditional Mean Process</th>
<th>Conditional Variance Process</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\beta_m$</td>
<td>$\beta_c$</td>
</tr>
<tr>
<td>AIRLN</td>
<td>0.318 (0.114)**</td>
<td>-0.261 (0.128)*</td>
</tr>
<tr>
<td>BANKS</td>
<td>0.038 (0.071)</td>
<td>0.107 (0.076)</td>
</tr>
<tr>
<td>ELECT</td>
<td>-0.049 (0.046)</td>
<td>0.136 (0.106)</td>
</tr>
<tr>
<td>GENIN</td>
<td>0.346 (0.038)**</td>
<td>-0.029 (0.032)</td>
</tr>
<tr>
<td>HOTEL</td>
<td>0.214 (0.102)*</td>
<td>0.121 (0.170)</td>
</tr>
<tr>
<td>MEDIA</td>
<td>0.112 (0.054)*</td>
<td>0.117 (0.064)</td>
</tr>
<tr>
<td>AUTOS</td>
<td>0.280 (0.072)**</td>
<td>-0.018 (0.092)</td>
</tr>
<tr>
<td>CHMCL</td>
<td>-0.027 (0.029)</td>
<td>-0.151 (0.034)**</td>
</tr>
<tr>
<td>ELETR</td>
<td>0.289 (0.082)**</td>
<td>0.095 (0.132)</td>
</tr>
<tr>
<td>ELTNC</td>
<td>0.514 (0.034)**</td>
<td>0.050 (0.026)</td>
</tr>
<tr>
<td>PAPER</td>
<td>0.268 (0.061)**</td>
<td>-0.279 (0.165)</td>
</tr>
<tr>
<td>PHARM</td>
<td>-0.264 (0.058)**</td>
<td>-0.111 (0.024)**</td>
</tr>
<tr>
<td>Factor risk premium</td>
<td>$m_0$</td>
<td>-0.003 (0.001)**</td>
</tr>
<tr>
<td></td>
<td>$m_1$</td>
<td>0.077 (0.001)**</td>
</tr>
</tbody>
</table>

Estimations are based on monthly dollar-denominated excess returns from 09/1978 through 12/2001. Each mean equation relates the excess industry returns $r_{it}$ to its world market and USD/JPY currency risks. The factor realizations, $F_{m,i}$ and $F_{c,i}$, depend on their own conditional volatility, and thus the expected factor risk premia, $\lambda_{m,i}$ and $\lambda_{c,i}$, are the functions of conditional volatility.

$r_{it} = (\lambda_{m,i} + \epsilon_{m,i})\beta_{im} + (\lambda_{c,i} + \epsilon_{c,i})(\beta_{ic} + \beta_d D_i) + \epsilon_{i,t} \forall i$

where $F_{m,i} = E(F_{m,i} | \Omega_{i-1}) + \epsilon_{m,i} = \lambda_{m,i} + \epsilon_{m,i} = m_0 + m_1 h_{m,i} + \epsilon_{m,i}$

$F_{c,i} = E(F_{c,i} | \Omega_{i-1}) + \epsilon_{c,i} = \lambda_{c,i} + \epsilon_{c,i} = c_0 + c_1 h_{c,i} + \epsilon_{c,i}$

$D_i$ is a dummy variable, which is equal to one if $\epsilon_{c,i} < 0$ and zero otherwise.

$\epsilon_i | \Omega_{i-1} \sim N(0,H_i)$

The conditional covariance matrix $H_t$ is parameterized as follows

$H_t = H_0 * (t^T - aa^T - bb^T) + aa^T * \epsilon_{i,t-1} \epsilon_{i,t-1}^T + bb^T * H_{t-1}$

where $H_t \in \mathbb{R}^{14x14}$ is the conditional covariance matrix of twelve industry returns and two risk factors. The elements of vectors $a, b \in \mathbb{R}^{14x1}$ are the GARCH parameters, $t$ is a 14 x 1 unit vector and * denotes the Hadamard product (element-by-element multiplication). QML standard errors are reported in parentheses. * and ** denote statistical significance at the 5% and 1% level, respectively.
### TABLE 6

<table>
<thead>
<tr>
<th>Factor risk exposure</th>
<th>Conditional Mean Process</th>
<th>Conditional Variance Process</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\beta_m$</td>
<td>$\beta_c$</td>
</tr>
<tr>
<td>BAC</td>
<td>0.008 (0.065)</td>
<td>0.089 (0.085)</td>
</tr>
<tr>
<td>WFC</td>
<td>0.033 (0.054)</td>
<td>0.130 (0.062)**</td>
</tr>
<tr>
<td>BK</td>
<td>-0.034 (0.058)</td>
<td>0.219 (0.076)**</td>
</tr>
<tr>
<td>MEL</td>
<td>0.004 (0.062)</td>
<td>0.076 (0.081)</td>
</tr>
<tr>
<td>NCC</td>
<td>0.008 (0.044)</td>
<td>0.093 (0.055)</td>
</tr>
<tr>
<td>PNC</td>
<td>0.044 (0.060)</td>
<td>0.107 (0.053)**</td>
</tr>
<tr>
<td>WB</td>
<td>0.033 (0.048)</td>
<td>0.143 (0.068)**</td>
</tr>
<tr>
<td>CMA</td>
<td>-0.349 (0.043)**</td>
<td>0.295 (0.058)**</td>
</tr>
<tr>
<td>NTRS</td>
<td>-0.019 (0.045)</td>
<td>0.174 (0.057)**</td>
</tr>
<tr>
<td>RIGS</td>
<td>0.234 (0.060)**</td>
<td>-0.235 (0.051)**</td>
</tr>
</tbody>
</table>

| Factor risk premium  | $m_0$ | 0.001 (0.000)** | 0.125 (0.011)** | 0.979 (0.005)** |
|                      | $m_1$ | 0.262 (0.042)** |
| USD/JPY              | $c_0$ | 0.001 (0.000) | 0.161 (0.012)** | 0.971 (0.004)** |

Estimations are based on weekly dollar-denominated excess US bank stock returns from 08/11/1978 through 12/28/01. Each mean equation relates the excess returns $r_{i,t}$ to its world market and USD/JPY currency risks. The factor realizations, $F_{m,t}$ and $F_{c,t}$, depend on its own conditional volatility, and thus the expected factor risk premia, $\lambda_{m,t}$ and $\lambda_{c,t}$, are the functions of conditional volatility.

$$r_{i,t} = (\lambda_{m,t} + \epsilon_{m,t})\beta_{m} + (\lambda_{c,t} + \epsilon_{c,t})(\beta_{c} + \beta^d_c D_t) + \epsilon_{i,t} \quad \forall i$$

where $F_{m,t} = E(F_{m,t} | \Omega_{t-1}) + \epsilon_{m,t} = \lambda_{m,t} + \epsilon_{m,t} = m_0 + m_1 h_{m,t} + \epsilon_{m,t}$,

$$F_{c,t} = E(F_{c,t} | \Omega_{t-1}) + \epsilon_{c,t} = \lambda_{c,t} + \epsilon_{c,t} = c_0 + c_1 h_{c,t} + \epsilon_{c,t}$$

$D_t$ is a dummy variable, which is equal to one if $\epsilon_{c,t} < 0$ and zero otherwise.

$$\epsilon_{i,t} | \Omega_{t-1} \sim N(0, H_i)$$

The conditional covariance matrix $H_t$ is parameterized as follows

$$H_t = H_0 * (u^T - aa^T - bb^T) + aa^T * \epsilon_{t-1} \epsilon_{t-1}^T + bb^T * H_{t-1}$$

where $H_t \in \mathbb{R}^{12 \times 12}$ is the conditional covariance matrix of ten bank stock returns and two risk factors. The elements of vectors $a, b \in \mathbb{R}^{12 \times 1}$ are the GARCH parameters, $t$ is a 12 x 1 unit vector and $*$ denotes the Hadamard product (element-by-element multiplication). QML standard errors are reported in parentheses. * and ** denote statistical significance at the 5% and 1% level, respectively.
### TABLE 7
Total Currency Exposures, the Degree of Asymmetry, and Summary Statistics of Risk Premia

<table>
<thead>
<tr>
<th></th>
<th>Panel A: Weekly USD/JPY</th>
<th>Panel B: Weekly TWFX</th>
<th>Panel C: Monthly USD/JPY</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Total currency exposures</td>
<td>Degree of asymmetry</td>
<td>Total currency exposures</td>
</tr>
<tr>
<td>AIRLN</td>
<td>0.033</td>
<td>1.997</td>
<td>0.174</td>
</tr>
<tr>
<td>BANKS</td>
<td>-0.05</td>
<td>-1.388</td>
<td>-0.066</td>
</tr>
<tr>
<td>ELECT</td>
<td>-0.034</td>
<td>-1.735</td>
<td>-0.119</td>
</tr>
<tr>
<td>GENIN</td>
<td>-0.041</td>
<td>-2.796</td>
<td>0.026</td>
</tr>
<tr>
<td>HOTEL</td>
<td>-0.019</td>
<td>-1.128</td>
<td>-0.002</td>
</tr>
<tr>
<td>MEDIA</td>
<td>-0.036</td>
<td>-1.530</td>
<td>-0.031</td>
</tr>
<tr>
<td>AUTOS</td>
<td>-0.027</td>
<td>-1.864</td>
<td>0.061</td>
</tr>
<tr>
<td>CHMCL</td>
<td>-0.051</td>
<td>-4.880</td>
<td>0.006</td>
</tr>
<tr>
<td>ELETR</td>
<td>-0.113</td>
<td>-2.240</td>
<td>0.057</td>
</tr>
<tr>
<td>ELTNC</td>
<td>-0.079</td>
<td>-6.201</td>
<td>0.017</td>
</tr>
<tr>
<td>PAPER</td>
<td>-0.01</td>
<td>-0.598</td>
<td>0.029</td>
</tr>
<tr>
<td>PHARM</td>
<td>-0.094</td>
<td>-1.793</td>
<td>-0.113</td>
</tr>
</tbody>
</table>

The total currency exposure is calculated by summing the estimated $\beta_c$ and $\beta_d$ coefficients. As for the degree of asymmetry, it is computed by taking the ratio $\beta_d / \beta_c$.  
