

**Asymmetrical Investor Sentiments in Country Funds:
The Effects of Domestic and Foreign Volume and Volatility on Fund Discounts**

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Abstract

Existing studies of country-fund premiums utilize market returns as main arguments and do not account for the effects of investor-risk perception specifically. This paper proposes a time-varying, asymmetrical investor-risk perception hypothesis that relates country-fund premiums or discounts to risk perceptions in U.S. and foreign stock markets. By using trading volume and conditional volatility to measure noise-trader risk (driven by sentiment) and rational-investor risk (motivated by fundamentals), respectively, we find significant evidence to support the time-varying, asymmetrical investor-risk hypothesis. Specifically, trading volume and volatility in both U.S. and foreign markets are crucial to explaining the behavioral changes of country-fund net premiums over time. While U.S. volume and volatility are more important than similar foreign variables on the whole, there exists significant regional heterogeneity, and foreign variables as well as exchange-rate volatility are important for a significant number of country funds. Fund premiums are also affected by the emerging-market financial crises, and there is some evidence of international contagion.

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1. Introduction

In the tradition of behavioral finance, Lee *et al.* (1991) suggest investor sentiment as an explanation of the closed-end fund puzzle. The persistent discounts that exist in closed-end funds defy arbitrage and thus reflect psychological-noise trading rather than economic fundamentals. Hardouvelis *et al.* (1994) note that the noise-trader model may be applicable to country funds, while Bodurtha *et al.* (1995) find that U.S. investor sentiment, in particular, can explain the discount on closed-end country funds traded in the United States. However, foreign local-market factors as well as U.S. market factors are found to be significant for country-fund pricing behavior in general [Choi and Lee (1996); Lee and Hong (2002)].

Closed-end country funds are traded in the United States, but are invested in a particular foreign country or region. Since frictionless international arbitrage could eliminate the divergence between the fund prices in the United States and the net asset values (NAV) of actual investment portfolios abroad, the investment barrier is one of the reasons that would explain why persistent divergence between the two might exist. Errunza *et al.* (1998) show such a possibility theoretically because of imperfect substitution of country funds and their underlying assets – that is, a country fund can trade at a premium because of barriers to international investment in partially segmented international capital markets. Bonser-Neal *et al.* (1990) provide empirical support for the view that relaxation of investment restrictions leads to lower fund premiums. Due to investment restrictions, the market-segmentation hypothesis may apply to a relatively small number of premium country funds but not to discount funds that are the absolute majority of closed-end country funds.

The literature on behavioral asset pricing suggests that asset returns in general reflect the sentiment of noise traders such as uninformed, small investors as well as the fundamental risk of rational investor-arbitrageurs. Dividing diverse investors into rational and irrational groups, DeLong *et al.* (1990), for instance, describe a market where rational investors form their expectations based on economic fundamentals, while the expectations of noise traders depend on their sentiments. In this setting, rational investors not only will have to manage risk on their actual portfolios, but also will have to assume the risk that noise traders may drive asset prices away from their fundamental values. Changes in noise-trader sentiment over time can be a source of risk for fundamental, risk-averse investors as well as for noise traders themselves. Thus the noise-trader risk is systematic and rewarded, and assets may be priced discounts relative to their fundamental values.

We argue that this phenomenon exists in closed-end, country-fund markets and that fund premiums or discounts can be explained by an interaction of fundamental investor risk and noise-trader sentiment in U.S. and foreign local markets. We include both U.S. and foreign local variables because the former affects fund prices while the latter influences net asset values. Fund premiums or discounts, in effect, represent asymmetrical, time-varying, investor-risk premiums – measuring asymmetrical investor sentiments by fundamental versus noise traders in the U.S. markets versus foreign local markets. Depending on the direction of these different investor risk and market sentiments, both premiums and discounts, *ex ante*, are possible. This contrasts with existing empirical work that utilizes *return* variables rather than specific risk measures and that focuses only on U.S. investor sentiment [e.g., Bodurtha *et al.* (1995)] or investment barriers [e.g., Bonser-Neal *et al.* (1990)]. The time-varying, asymmetrical investor-risk perception view

presumes partial segmentation as an international market environment but does not depend on it for explanation of premiums or discounts.¹

We focus on trading volume and market volatility as parsimonious measures of noise-trader sentiment and fundamental-investor risk, respectively. Market volatility is a proxy for investor assessment of fundamental risks stemming from informational trading, while trading volumes signify noise-trader risk perceptions resulting from non-informational, psychological trading [Campbell *et al.* (1993); Shalen (1993); Lee and Rui (2001); Hirshleifer and Teoh (2003)]. If the fund premium or discount is a proxy for some unspecified time-varying risk premium, the net fund premium should be related to trading volume that reflects noise-trader risk as well as market volatility that reflects fundamental risk.

Based on 22 closed-end, single-country funds, we find that U.S. trading volume and market volatility are more important than similar foreign variables. However, significant regional differences exist, and premiums depend on foreign volume and volatility (and exchange-rate risk) for a significant number of country funds. Fund premiums also vary temporally with crises – emerging-market financial crises tend to create discounts on country funds, although funds that different characteristics of investors are measured, respectively, by market volatility and trading volume in both the U.S. and foreign markets.

The rest of the paper is organized as follows. Section 2 presents empirical models and hypotheses. Section 3 discusses the data and the measurements of the variables. Section 4 reports

¹ An alternative explanation for fund premiums and discounts – that is still consistent with market segmentation – is differential investor sophistication. Seasholes (2004), Albuquerque, et al. (2005) and others argue that U.S. investors may be more sophisticated than foreign investors in security valuation. If this is the case in country funds, both fund premiums and discounts are possible because U.S. investors may value underlying assets at a different level than foreign local investors. However, most investors of country funds in the U.S. tend to be individual investors rather than institutional investors, while most foreign securities that country funds own (i.e., large firms) also tend to be owned by many institutional

empirical results of the basic model regarding the effects of volume and volatility on net fund premiums. Sector 5 examines the impact of the emerging-market financial crises. Section 6 extends the model to include return measures in addition to risk variables, and estimates fund returns and NAV returns in addition to fund premiums. Section 7 summarizes our conclusions.

2. Empirical Models and Hypotheses

Lee *et al.* (1991) contend that the traditional argument for dual country-fund behavior – market segmentation due to taxes, investment barriers, or country risks [e.g., Errunza *et al.* (1998); Bonser-Neal *et al.* (1990)] – is insufficient to explain the dynamic behavior of country-fund discounts. The existence of noise traders may create additional risk for rational investors because changes in noise-trader sentiments may create misperceptions about fundamental market judgments by rational traders. If noise-trader risk is systematic, the resulting risk premium implies that fund prices are determined at a discount relative to their fundamental values [DeLong *et al.* (1990)].²

Existing empirical studies of country funds [Bodurtha *et al.* (1995); Chang *et al.* (1995); Arshanapalli *et al.* (1996); Choi and Lee (1996); Lee and Hong (2002)] utilize multifactor models where fund returns are regressed on U.S. stock market returns and local stock-market returns (and exchange-rate changes).³ The use of market-return measures, however, may not

investors. Therefore, the notion of greater sophistication of U.S. investors than foreign investors may be less convincing in the context of country funds.

² An exposition of a noise-trader model is found in Shleifer and Summers (1990). Studies such as Neal and Wheatley (1998), Gemmill and Thomas (2002), and Lee *et al.* (2002) show that domestic closed-end fund discounts are negatively related to noise-investor sentiment, while Elton *et al.* (1998) and Sias *et al.* (2001) indicate that noise-trader risk is not priced.

³ This is at odds with the general international asset-pricing literature [survey by Stulz (1994)] where local and global market returns are included. However, given the particular duality of country funds – funds are traded in the United States, while the investments are made in the local market – the popular specification

directly capture the effects of investor sentiments or risk perceptions, which may be important to explain fund discounts. Neal (1994, p. 400) notes, “It is possible that country fund sentiment is a proxy for some unspecified time-varying risk premium.” Similarly, Hardouvelis *et al.* (1994, p. 386) state: “If markets are segmented, innovations in the ratio of the domestic price of risk to the foreign market price of risk can affect the discounts.” If the fund discount or premium is a proxy for some unspecified time-varying risk premium, the net fund premium should be related to trading volume that reflects noise-trader risk as well as to market volatility that reflects the fundamental risk of rational investors.

Lee and Rui (2001) show that trading volume reacts to the nature of new information. Homogenous public information that affects all investors is not likely to generate large trading volume. However, interactions among different groups of investors with heterogeneous information, such as selling pressure, have a significant effect on trading volume. Klibanoff *et al.* (1998) show that volume is higher during a news week as irrational investors increase their participation in the market. Campbell *et al.* (1993) report that uninformed selling trades by noise traders accompanied by high trading volume have a low serial correlation in returns because rational investors would require higher expected return for bearing noise-trader risk. Cooper (1999) and Kelly (1997) also report that noise variables increase trading volume.

In this study, we contend that country-fund discounts are explained by asymmetrical investor risk perceptions measured by trading volumes and market volatilities. Country funds are subject to noise-trader risk perceptions (that can be measured by trading volume) as well as to fundamental market volatility faced by rational investors. However, since fund shares are priced in the United States, while the net asset values are determined in local markets, we hypothesize

is sensible from the U.S.-investor perspective. This method amounts to an assumption that U.S. and world

that fund premiums are related to the conditional variances of aggregate stock-market returns and trading volumes in both U.S. and local markets. Our basic empirical model for net fund premium, $PR_{i,t}$, is as follows:

$$PR_{i,t} = \alpha + \gamma_m V_{m,t} + \gamma_m^* V_{m,t}^* + \lambda_m \sigma_{m,t}^{e,2} + \lambda_m^* \sigma_{m,t,LC}^{e*,2} + \delta_s \sigma_{s,t}^{e,2} + \varepsilon_{i,t}, \quad (1)$$

where $PR_{i,t}$ is defined as the natural log-difference between country-fund i 's price and its NAV in U.S. dollar terms at time t ; $V_{m,t}$ is the detrended trading volume in the U.S. stock market; $V_{m,t}^*$ is the detrended trading volume in the corresponding foreign local stock market; $\sigma_{m,t}^{e,2}$ is the expected volatility of U.S. market returns; $\sigma_{m,t,LC}^{e*,2}$ is the expected volatility of foreign stock-market returns in local currency; and $\sigma_{s,t}^{e,2}$ is the expected volatility of changes in exchange rates.⁴ Because of the concern on contemporaneous correlations between volatility and trading volumes [Lamoureux and Lastrapes (1990); Shalen (1993)], we have orthogonalized the expected volatilities of returns and trading volumes in each market by regressing conditional volatilities on trading volumes and taking the residuals as conditional volatility measures in Equation (1).⁵

Relating trading volume and noise trading more specifically, it is arguable that high trading volume reflects both optimistic and pessimistic sentiments by irrational investors. When noise traders are irrationally optimistic, extra demand for shares is created, driving prices up and

markets are highly correlated.

⁴ The exchange-rate volatility term separates exchange-rate risk from foreign-stock market risk. The interaction between the two markets is not considered in this paper.

⁵ In lieu of market volatilities and volumes, similar variables for individual funds can be used as explanatory variables. However, individual fund variables measure asset-specific risks rather than market risks. We use market volatilities and volumes because in this paper we are interested in capturing the

increasing fund premiums (or reducing discounts). When investors are irrationally pessimistic, an increase in supply will lower prices and fund premiums (or increase fund discounts). Thus, it looks as if the relationship between trading volume and fund premiums were mixed. However, an increase in trading volume, by itself, means a higher level of noise-trader risk, which is borne by both rational and irrational investors and thus may be systematic. An arbitrageur who buys fund shares and sells the underlying shares short faces resale-price risk. If the arbitrageur has to liquidate a position in finite time, the risk from the unpredictability of noise-trader sentiment at the time of liquidation prevents aggressive trading that would eliminate discounts. Moreover, since noise-trader risk exists when small investors become pessimistic, not when they are optimistic, the case of fund-premium decrease (or fund-discount increase) will dominate the case of fund-premium increase (or fund-discount decrease). This asymmetrical investor-risk behavior indicates that, on average, an increase in trading volume in the U.S. stock market is associated with a decline in fund price, leading to an expected fund discount relative to the fund's fundamental values, that is, $\gamma_m < 0$.

Similar logic can apply to market volatility, only this time in terms of rational investors motivated by market fundamentals. To the extent that an increase in conditional U.S. market volatility requires an extra risk premium, we would expect $\lambda_m < 0$.⁶ With an increase in conditional volatility, a higher risk premium by rational risk-averse investors will require lower fund prices and hence a discount on fund prices. Conditional variance provides broad information to summarize the risk present in the market; in the GARCH specification, the lagged

effects of risk perceptions at the market level in this paper. Preliminary estimations indicate that market volatility and volume measures have somewhat greater explanatory powers than similar fund variables.

⁶ For a positive relationship between asset return and volatility, see French *et al.* (1987) and Campbell and Hentschel (1992). However, some other studies, including Breen, Glosten and Jagannathan (1989) and Turner, Startz and Nelson (1989), report a negative relation.

dependent variable summarizes all the lagged information associated with the lagged independent variables.⁷

With respect to foreign-market volume and volatility, we anticipate that $\gamma_m^* > 0$ and $\lambda_m^* > 0$ in principle. This is because an increase in foreign trading volume or volatility, which represents an increase in irrational and fundamental risk perceptions in the foreign market respectively, will reduce the NAV and, therefore, increase fund premiums. However, it is also possible that country-fund premiums reflect the effects of international contagion [e.g., Frankel and Schmukler (1996)]. Errunza *et al.* (1998) show the existence of noise-spillover effect. If so, a rise in the perceived risk of the foreign stock markets, measured by foreign-trading volume and conditional market volatility, will induce U.S. investors to demand a higher return on country funds to compensate for the extra risk. This will lower the fund price relative to the corresponding NAV, producing a reduction in premiums. Thus, the signs of γ_m^* and λ_m^* can be mixed, depending on the relative force of the positive within-market risk effects versus the negative international-contagion effects.

Finally, the effect of the conditional variance of changes in exchange rates is expected to be positive. When investors expect increased exchange-rate volatility, it becomes more desirable to purchase country funds than foreign stocks denominated in the foreign currency. Since country funds are traded in the U.S. dollar, they provide a means of avoiding costly exchange-risk hedging. As a result, demand for country funds may increase relative to foreign stocks, bidding up the prices of country funds. Thus we then expect country-fund premiums to be positively affected by the expected volatility in the currency markets, i.e., $\delta_s > 0$.

⁷ Recently, Goyal and Santa-Clara (2003) find some evidence that even an idiosyncratic risk may be priced in stock markets.

3. Data and Descriptive Statistics

The sample in this study covers the period from January 2, 1987, to May 26, 2000. The sample period spans more than twelve years and includes a period of major emerging-market financial crises. Table 1 lists the 22 funds included in the sample, along with their initial public offering (IPO) dates, sample starting dates, and the number of observations. If there are multiple funds from a given country, we include only one fund from that country – the one with the longest history – because of the commonality of market variables for funds investing in that country. The 22 closed-end country funds are classified in terms of the regions or the level of economic development of the host country. There are three Latin-American funds, twelve Asian-Pacific funds, and seven European funds. Alternatively, following the criteria advanced by Morgan Stanley Capital International, there are twelve emerging-market funds and ten developed-market funds.

For each fund, we obtain weekly fund prices and net asset values (NAVs) from *Lipper Analytical Services*, adjusted for splits, right offerings, and in-kind distributions. Both fund prices and NAVs are reported in U.S. dollars as of Friday's close.⁸ Since certain foreign stock markets may close several hours prior to the close of New York City markets, the reported fund prices and NAVs are only approximately synchronous, even though they correspond to the same calendar-day trading sessions. The bias arising from imperfect synchronous trading can be potentially serious in high-frequency, short-horizon data such as daily data, but less so in weekly

⁸ Several funds in our sample do not report their NAVs as of the Friday close. The India Growth Fund reports its NAV as of the Wednesday close, and Mexico, Taiwan, and Singapore funds report their NAVs as of the Thursday close. For these funds, we use fund prices and exchange rates on the day the NAVs are reported.

data. To address this concern, we followed Bodurtha *et al.* (1995) and experiment with lagged data, although no significant differences were found.

We obtain weekly foreign stock-market indexes from *Datastream*, and foreign exchange rates (measured by the U.S. dollar per unit of foreign currency) from the *Federal Reserve Bank of Chicago* and *Datastream*.⁹ The foreign exchange-rates data from the *Federal Reserve Bank of Chicago* are noon buying rates in New York City and those from *Datastream* are as of the close of business in the country of origin. Table 1 lists the specific foreign-market indexes used in this study. We also collect from *Datastream* the weekly Standard and Poor's 500 stock market index for the United States.

Weekly transaction volume in values and number of shares are collected from *Datastream* for the U.S. and foreign stock markets. We use weekly turnover in value terms as the measure of the trading volume adjusted for linear and nonlinear time trends [Gallant *et al.* (1992); Lee and Rui (2001)]. Turnover is defined as the ratio of the total market value of shares traded during the day of NAV reporting to the total market value of shares outstanding. Turnover is widely used in existing work as a measure of volume and has some advantages. First, the use of turnover helps to reduce low-frequency variation in time series because both the number of shares outstanding and the number of shares traded have grown steadily over time [Campbell and Kyle (1993)]. Second, it facilitates comparison of trading activities across different markets and securities since the turnover is unit-free and a standardized measure [Freund and Webb (1999)].

Conditional volatilities in U.S. and foreign markets, used as data in Equation (1), are estimated by a multivariate GARCH(1,1) model, with U.S. stock market returns, foreign stock

⁹ *Datastream* is used when exchange rates are not available from the *Federal Reserve Bank of Chicago*. Such currencies include those of Indonesia, the Philippines, Turkey, Argentina, Chile, and Mexico.

market returns in local currency, and foreign exchange returns.¹⁰ The standard GARCH(1,1) specification provides an excellent fit for the time-varying volatility of stock returns [Bollerslev *et al.* (1992); Franses and van Dijk (1996)]. To guarantee a positive definite variance/covariance matrix, the parameterization procedure suggested by Engle and Kroner (1995) is used, and the quasi-maximum likelihood estimation method proposed by Bollerslev and Wooldridge (1992) is used to facilitate inference in the case when the conditional distribution of the residuals is non-normal.

Table 1 presents means and standard deviations of country-fund premiums or discounts, country-fund returns, and NAV returns, respectively. Fund premiums or discounts are significantly different from zero for all funds but one (Turkish Investment Fund). Most funds are traded on average discounts, except for Asian country funds such as Indonesia, Korea, Malaysia, Taiwan, Thai, and Japan OTC equity funds. Those funds are from countries severely affected by the Asian financial crisis and display average fund premiums for the period. These results are consistent with Chandar and Patro (2000), who report that Asian funds have experienced a significant increase in premiums during the crisis. Average fund returns and NAV returns are generally statistically significant, but positive for European and Latin American funds and negative for Asian-Pacific funds.

Statistics on the second moment show that fund returns and NAV returns for emerging markets appear to be more volatile than those for developed markets. As for the fund premiums, eight out of the ten most volatile funds are Asian-Pacific emerging-market funds. In addition,

¹⁰ Since conditional variances are estimated from a model rather than data, they may be open to a potential “stochastic regressor” problem. However, Pagan (1984) shows that if the first step produces consistent estimates of true conditional variance and covariance, the second step will produce consistent estimates of the parameters of interest. Moreover, the GARCH specification used to estimate conditional

consistent with previous studies [e.g., Bodurtha *et al.* (1995)], fund returns are more volatile than NAV returns over time.

4. Estimation of the Basic Model

A. Estimation Results

Since all country funds in the sample are traded in New York City and subject to common shocks in the U.S. market, the error terms of the estimated equations are likely to be correlated across the funds. For this reason, the fund-premium equations are estimated jointly by the seemingly unrelated regression (SUR) method proposed by Zellner (1962). Also, because of concern about potential serial correlation and heteroskedasticity problems in residuals, the Newey-West (1987) method is used to achieve consistent estimates. For the purpose of SUR estimations, all country funds are divided into four regional groups (Latin American, emerging Asian-Pacific, developed Asian-Pacific, and European groups), and estimations are conducted jointly within each group.¹¹

Table 2 presents the estimation results of Equation (1) for all 22 country funds. Except for two funds (Taiwan and Thai), the coefficients of the U.S.-market trading volume have negative signs as expected. The U.S. volume coefficients of all Latin American and European country funds are also statistically significant at the one percent level. However, only four (China, India Growth, Korea, and First Australia) out of twelve Asia-Pacific funds show significance at even

variances in the first step is consistent with an instrumental variable approach when a lagged dependent variable is used as an instrument.

¹¹ The SUR procedure requires all funds to have the same data period. The starting dates of each group are thus decided by the funds that have the shortest data period: October 25, 1991, for the Latin-American funds (set by Argentina Fund); July 10, 1992, for emerging Asian-Pacific funds (set by China Fund); August 2, 1990, for developed Asian-Pacific funds (set by Singapore Fund); and May 11, 1990, for European funds (set by France Growth Fund). The ending date is May 26, 2000, for all fund groups.

the ten percent level. As we will examine separately, this may be related to the effects of the Asian financial crisis. In general, the evidence supports the hypothesis that country-fund premiums are negatively correlated with U.S. aggregate stock-trading volumes. This is consistent with the market behavior in which an increased risk perception of noise traders leads to an increase in trading volumes, which in turn leads to a decrease in fund prices and causes an increase in fund discounts (or a decrease in premiums).

On the other side of the market, the coefficients of foreign-market trading volumes show mixed signs on fund premiums – twelve funds have negative signs and ten funds have positive signs. This is consistent with the presence of the contagion effect. An increase in foreign trading volume would reduce the NAV and improve fund premiums. However, this is offset by the negative impact of international contagion on fund prices and premiums in the United States. Because of the conflict, the net effects are generally weak. Of the 22 total funds, only nine show statistically significant coefficients of foreign-market trading volumes on the fund premiums, of which five are negative and four are positive.

Combining domestic and foreign volume effects, we can conclude that the volume effects on fund premiums, in principle, are negative. However, in the case of foreign volume, the signs are split because of the influence of foreign contagion effects on U.S. fund prices. It is also noteworthy that the effects of U.S. volume are most pronounced for European and Latin-American funds, while the effects of foreign volume are most significant for the Asian funds. Again, as we will examine separately, we think it is related to the effects of the emerging-market financial crises.

Alternatively, all 22 funds can be estimated jointly by the SUR; however, this would be computationally inefficient.

Table 2 also presents the effect of domestic-market and foreign-market volatility. Regarding the effects of domestic volatility in the U.S. market, the theoretical section above predicts the coefficient to be negative. This is consistent with the notion that, if investors perceive higher volatility in the U.S. market, they will require higher expected returns, which is possible when fund prices decline. Such prediction is borne out for Asian and Latin-American funds because twelve out of thirteen significant coefficients (out of the total of fifteen such funds) are negative. Three European funds (Germany, Spain, and Swiss Helvetia), however, show positive coefficients. This result for European funds is rather odd. One conjecture is that, to the extent that the European markets are more correlated with the United States than with emerging markets, the high U.S. market volatility may cause U.S. investors to reallocate their portfolios into the European funds, bidding up the prices of those European funds.

With respect to the effect of the conditional variance of foreign stock-market returns, only nine coefficients are statistically significant. As in the case of foreign volume, the signs are also mixed. Of the nine significant coefficients, six positive ones are all for emerging-market funds, while three negative ones are for European funds. The negative signs are as expected from the effects of investor-risk premiums, but the positive effect of volatility may be attributable to the uncertainty stemming from emerging-market financial crises. High volatility in foreign stock markets during the crises can create fund premiums because fund prices in the United States may not fall as much as the decline in NAVs in foreign markets.

Finally, fund premiums in U.S. dollar terms may respond to exchange-rate changes as well as to investor-risk perceptions in the U.S. and foreign stock markets. If so, the variability of fund premiums would reflect conditional variance-covariance to changes in stocks and exchange rates. The variance of exchange-rate changes is thus included in Equation (1). The results are

mixed, as only ten coefficients out of 22 are significant. However, it is interesting that all ten significant exchange-volatility coefficients are for Asia-Pacific and Latin American funds. Not a single European fund has a significant exchange-volatility coefficient. Again, this may reflect the role of currency volatility in Asian and emerging-market crises.¹² Note that, except for two funds (First Australia and Taiwan), the signs of all significant coefficients for emerging-market funds are positive. The positive association of fund premiums and expected exchange risk underscores the basically speculative nature of these markets – an anticipation of higher exchange-rate volatility creates an option-like value because of the potential for profit during a financial crisis. Alternatively, it is also consistent with a notion that exchange risk can be more easily hedged against European currencies than against emerging-market currencies.

In sum, as it pertains to the impact on fund premiums, U.S.-market volume and volatility is more important than similar foreign variables. At the ten percent level of significance, fourteen and sixteen funds, respectively, show significant coefficients for U.S. volume and volatility, in contrast to only nine funds for both foreign volume and volatility. Exchange-rate volatility shows significance for ten funds. The greater significance of U.S. variables appears to be consistent with the U.S.-sentiment hypothesis advanced by Bodurtha *et al.* (1995). However, as shown in the case of nine funds, the foreign-market effects are also important for a sizable number of funds. More importantly, in contrast to existing work that depends on *return* variables, we focus on the impact of volume and volatility as measures of investor risks perceived by rational and irrational investors. In addition, the effects of foreign variables may differ significantly by region. Foreign-trading volumes have significant effects on premiums for Asian and Latin-

¹² Bailey *et al.* (2000) examine the effect of exchange-rate news on emerging-market country funds during the 1994-95 Mexican financial crises.

American funds, but not for European funds, while foreign-market volatility positively affects Asian and Latin American emerging-market funds and negatively affects the European funds. These differences appear to be related to the impact of emerging-market financial crises occurring during the data period.

B. Test of Significance across U.S. and Local Markets

We now examine formally whether significant differences exist in the effects of U.S. versus foreign-market variables and also across the regions. Specifically, we test (a) whether the coefficients of U.S. variables (trading volume and volatility) are significantly different from those of foreign variables and (b) whether the coefficients of such variables are significantly different across regions.

The χ^2 statistics in Table 3, Panel A, indicate that the null hypotheses of identical U.S. and foreign volume effects are rejected for most of the funds (twelve out of 22 funds for volume and sixteen for volatility). Considering volume and volatility effects combined, U.S. and foreign effects appear to differ significantly for most cases. This is at odds with Bodurtha *et al.* (1995) who report the “U.S. sentiment” as a sole explanation of fund premiums.

In Panel B, we test the null hypothesis of the equality of coefficients for each risk factor across funds. The results for all funds together show that the null hypothesis is rejected, implying that different funds have different sensitivity to the U.S. and foreign risk factors. The similar results – rejection of the null of equal coefficients – are obtained for all risk variables for emerging-market or develop-country funds. Regionally, the null is rejected for most of the groups and for most variables. In sum, these tests suggest that, despite their commonality as closed-end country funds traded in the United States, the sensitivity of fund premiums to risk factors can be diverse, depending on local-market characteristics and country of origin.

5. Impact of Emerging-Market Crises

The estimations obtained above may be subject to structural bias due to major emerging-market financial crises, such as the Mexican peso crisis in December 1994 and the Asian financial crisis that began in July 1997. Frankel and Schmukler (1996) argue that since local investors have superior information about their own markets compared with foreign (U.S.) investors, a discount is required to entice U.S. investors to hold the foreign country funds. Their study of three Mexican country funds during the 1994 Mexican peso crisis provides evidence in support of an asymmetrical-information hypothesis.¹³ Chandar and Patro (2000) find that country-fund prices are more sensitive to global-market elements, while the NAVs are influenced more significantly by local-market conditions and that this difference in risk exposure has been magnified during the crisis. However, these studies concentrate on returns and do not utilize risk measures despite the turmoil and volatility caused by the crises.

To incorporate the impact of financial crises, an intercept dummy is included in Equation (1). We include a zero-one dummy for the Asian crisis for the period from July 4, 1997, to May 26, 2000, and another one for the Mexican peso crisis for the period from December 2, 1994, to December 29, 1995. These dates cover the period from the beginning of the crises and periods of ensuing turmoil.

The coefficients of the dummy variables reported in Table 4 are significant for most fund premiums. The Asian-crisis dummies are significant for eighteen funds, and the Mexican fund

¹³ The asymmetrical information hypothesis of Frankel and Schmukler (1996) differs from the noise-trader model that we use. Information asymmetry is assumed between the local and foreign investors in the former model, while the different risk perceptions are defined for rational and irrational investors in each market in the latter.

dummies for thirteen funds. Thus, the impact of the Asian financial crisis is shown to be more pervasive than the Mexican-peso crisis. It is also interesting that the impact of the crises transcends the particular region where the crisis began, or even emerging markets as a whole. Many European funds are affected by these crises emanating from Asian or Latin-American emerging-market countries. This suggests the role of international contagion, or an investment behavior that regards all country funds as more or less a single-asset category and reacts accordingly to external shocks.

Generally, one would expect the effects of these crises on country-fund prices to be negative. And that is in fact the case for all European funds. However, if we look at the Asian funds that presumably are most directly affected during the Asian crisis, four Asian funds (Indonesia, Thai, Malaysia, and First Philippines) out of nine show statistical significance and have positive coefficients rather than negative ones. The reason is that declines in NAVs in the local markets are so precipitous as to outweigh the decline in fund prices in the United States, creating fund premiums.¹⁴ Similarly, during the Mexican-peso crisis, the effect on Mexico Fund is significantly positive, while it is negative for most of the other funds, including all European funds.

Finally, Panel B of Table 4 shows that the null hypotheses of identical coefficients across funds are rejected for all categories at least at the five percent level of significance. This indicates that the impact of crises on net premiums vary from fund to fund. These results suggest that while the impact of major emerging-market financial crises are often traumatic and pervasive, the premiums of the affected funds can be positive, as the price declines might be

¹⁴ Of the Asian countries that are directly affected by the Asian financial crisis, the Korea fund is an exception because it shows a significant negative crisis dummy. This can be attributable to the rapid

more severe in the U.S. market than in the local markets. This finding supports the time-varying asymmetrical risk-perception hypothesis and is also consistent with Frankel and Schmukler (1996) and Chandar and Patro (2000). However, Table 4 also indicates that, despite the generally significant impact of the crises, the degree to which a particular country fund is affected is sufficiently diverse.

6. Inclusion of Returns and Estimation of Fund Returns and NAV Returns

The basic model in Section 4 is now extended in two respects. First, for purposes of comparison with existing work (e.g., Bodurtha et al., 1995; Choi and Lee, 1996; Lee and Hong, 2002) that utilize domestic and foreign returns, we now include U.S. and foreign returns and exchange-rate changes in addition to domestic and foreign volume and volatility in the fund premium equation. Second, since fund premiums are the logarithmic difference between fund prices and NAVs, it raises the question of how these component variables behave to arrive at a fund premium or discount. To address this concern, we estimate the fund returns and NAV returns in addition to net fund premiums.

A. Inclusion of Returns

Bodurtha *et al.* (1995) study the behavior of country-fund net premiums, focusing on U.S. market sentiment. They show that premiums of individual country funds tend to move together because of the co-movement of their stock prices with the U.S. market. An empirical equation for net fund premiums used by Bodurtha *et al.* (1995) is:

$$PR_{i,t} = \alpha + \sum_{i=-1}^1 \beta_i^* R_{m,t+i}^* + \sum_{i=-1}^1 \phi_i \Delta S_{t+i} + \beta_i R_{m,t} + \varepsilon_{i,t}, \quad (2)$$

recovery and subsequent slowdown of the Korean economy during the Asian crisis dummy period of July 1997 – May 2000.

where $R_{m,t}^*$ is foreign stock-market returns in the local currency unit at time t ; ΔS_t is $\log(S_t/S_{t-1})$ where S_t is the U.S. dollar per unit of the foreign currency at time t ; $R_{m,t}$ is U.S. stock market returns in U.S. dollars at time t ; α , β^* , ϕ , and β are parameters; and $\varepsilon_{i,t}$ is a random-error term.¹⁵ One-period leads and lags are included for foreign returns to adjust for the effects of time differences between U.S. and foreign markets.

We now estimate an augmented model which combines Equation (1) and Equation (2):

$$PR_{i,t} = \alpha + \gamma_m V_{m,t} + \gamma_m^* V_{m,t}^* + \lambda_m \sigma_{m,t}^{e,2} + \lambda_m^* \sigma_{m,t,LC}^{e*,2} + \delta_s \sigma_{s,t}^{e,2} + \sum_{i=-1}^1 \beta_i^* R_{m,t+i}^* + \sum_{i=-1}^1 \phi_i \Delta S_{t+i} + \beta_i R_{m,t} + \varepsilon_{i,t} \quad (3)$$

where U.S. stock-market returns, foreign stock-market returns and foreign-exchange returns are included as per Bodurtha *et al.* (1995) in addition to variables of volume and volatility. Table 5 reports the results.

The first thing to notice in Table 5 is that, compared to our basic model (Table 2), the overall improvement in R -squared terms is rather limited, suggesting that return variables may not be adding significantly to the explanatory power of country-fund premiums. In addition, the overall pattern of the coefficients of individual risk factors is roughly the same as those obtained in Table 2. The coefficients of the U.S. trading volume are generally negative and statistically significant, especially for the European and Latin-American funds, supporting the noise-trader risk hypothesis. Similarly, the effects of U.S. and foreign market volatility are the same, with exactly the same number of significant coefficients as before (sixteen and nine, respectively). However, foreign trading volume has only six significant coefficients compared with nine in

¹⁵ Bodurtha *et al.* (1995) actually use the *changes* in country-fund premiums as the dependent variable and include an equally weighted index of premium changes on domestic closed-end funds in the United States as another explanatory variable. We did not follow this exactly because fund returns are stationary

Table 2. The effect of exchange-rate volatility also drops slightly, with eight significant coefficients now as opposed to ten in the basic model.

With respect to the coefficients of return variables specifically, only two out of 22 funds show significance for the U.S.-market variable. This result is not consistent with the evidence reported by Bodurtha *et al.* (1995) that indicates the significance of U.S.-market returns. Although it is possible that market sentiment can be captured by volume and volatility rather than the market return as such, the present results show that the effects of the U.S.-market returns diminish in the augmented model, and, as we will see below, the foreign-market returns may also have additional influences, supporting the findings of Choi and Lee (1996) and Lee and Hong (2002).

Several findings are worth noting with regard to the foreign-market coefficients. First, the lagged coefficients are positive and significant for most of the Asian-Pacific funds, indicating that the previous week's foreign-stock return has some information content in predicting current fund premiums in the U.S. market. Given the weekly data, any bias stemming from nonsynchronous market openings pales compared with the effect due to information content. Second, the majority of coefficients of European-market returns (and all four significant coefficients out of seven) are negative. That is, a decrease in fund premiums in the U.S. market is associated with an increase in the European market. This implies that when European stock markets rise, American investors can gain from investing there directly rather than from buying European country funds in the United States. Third, the coefficients of the lagged foreign-exchange returns are positive and significant for five out of seven European funds, indicating that a depreciation of the U.S. dollar (or an appreciation of the European currencies) leads to a

and thus taking changes in returns may introduce instability and therefore may not be appropriate for

rise in current fund premiums. This suggests a tendency that American investors invest in the European-country funds when they expect strong European currencies.

B. Estimation of Fund Returns and NAV Returns

Since fund premiums are the logarithmic difference between fund prices and NAVs, there is a question of how these component variables behave in addition to fund premiums. To address this concern, we now estimate fund returns and NAV returns using the same set of explanatory variables that include returns, volume, and volatility. A summary of estimation results is presented in Table 6.

Table 6 shows that domestic U.S.-market returns are significant for almost all of the fund returns, although less so for NAV returns as well as fund premiums. This is consistent with Bodurtha *et al.* (1995) who find that country-fund returns are dominantly influenced by U.S.-market returns rather than by NAV returns. However, foreign-market returns are also significant for almost all fund returns and NAV returns as well. Therefore, the coefficients on country-fund returns are generally positively correlated with both U.S.-market and foreign-market returns, supporting the dual characteristics of country funds [Choi and Lee (1996); Lee and Hong (2002)] rather than the U.S. sentiment hypothesis alone. Moreover, fund returns show a positive relation with foreign-exchange returns, which is generally significant, especially for Latin-American and European funds.

Another interesting result of Table 6 is that the level of significance of risk factors – volume and volatility – is much smaller in the fund return and NAV return equations than in the fund premiums. However, regardless of whether they are of domestic or foreign origin, the effects are generally greater for NAV returns than for fund returns. Overall there is an interesting

ordinary regressions [Arshanapalli *et al.* (1996); Chiang and Kim (2003)].

dichotomy – return measures are generally more significant for fund returns and NAV returns than for fund premiums, while volume and volatility effects are greater for fund premiums than fund returns and NAV returns. This underscores the importance of risk measures for fund premiums and the fact that the addition of return variables does not change the conclusion obtained from the basic model with regard to the impact of volume and volatility on fund premiums. The present paper therefore provides consistent evidence, based on risk-perception measures of noise and rational investors, in support of the time-varying, asymmetrical risk-perception hypothesis in both U.S. and local markets.

7. Summary and Conclusions

Existing studies of country funds such as Hardouvelis *et al.* (1994) and Bodurtha *et al.* (1995) emphasize the importance of U.S. investor sentiment, but utilize return measures and do not consider risk measures that characterize specific investor behaviors. Choi and Lee (1996) and Lee and Hong (2002) show that foreign-market factors as well as domestic U.S. factors are important. In this paper, we estimate the effects of specific market-risk measures such as volume and volatility in both U.S. and foreign markets on country-fund premiums. We argue that mainly uninformed irrational traders drive trading volumes, on the margin, while expected market volatilities are triggered primarily by activities of informed rational traders. Thus, volume and volatility measure noise-trader risk (driven by sentiment) and rational-investor risk (motivated by fundamentals), respectively.

Specifically, we propose a time-varying, asymmetrical investor-risk perception hypothesis that relates country-fund premiums or discounts to these risk factors in the United States and foreign stock markets. Fund premiums and discounts result from time-varying interactions of

asymmetrical rational and irrational investor-risk perceptions in the U.S. and foreign local markets. Using 22 U.S.-traded, closed-end, single-country funds, we find some significant evidence to support such a hypothesis. While U.S. volume and volatility variables are clearly more important than similar foreign variables on the whole, the patterns are quite heterogeneous. For instance, the U.S. volume variables are significant for European funds, while foreign volume as well as the U.S. and foreign market volatilities including exchange rate volatility significantly affect Asian funds. The results are robust when returns as well as risk measures are included in the model. It is also interesting that country-fund premiums are more strongly influenced by risk factors than are fund returns or NAV returns. In addition, fund premiums are significantly affected by the Asian financial crisis and, to a somewhat lesser extent, by the Mexican peso crisis. However, country funds that are directly affected by the crisis actually experience fund premiums rather than discounts because foreign local investors may react more negatively towards NAVs than U.S. investors do towards fund prices in the United States.

In sum, we conclude that country-fund net premiums may be sensitive to risk perceptions of irrational and rational investors in both U.S. and local markets. While U.S. risk variables are generally more important than similar foreign variables, there is evidence that significant regional differences exist and that foreign volume and volatility variables, including exchange-rate volatility, are important for a significant number of funds. As such, it underscores the asymmetrical role of U.S. and foreign noise and fundamental investors in determining country-fund premiums. These results are consistent with the time-varying, asymmetrical, risk-perception hypothesis and extend existing work that focuses on U.S. investor sentiment or that utilizes return measures rather than risks specifically.

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Table 1
Descriptive Statistics on Closed-End Country Funds

Fund	IPO Date	Sample Date	N	Foreign Stock Index	Fund Premium		Fund Returns		NAV Returns	
					Mean	S.D.	Mean	S.D.	Mean	S.D.
A. Latin-American Funds: Emerging Market										
Argentina	911011	911025	448	Data Stream Price Index	-0.0603***	0.1650	-0.0010	0.0477	0.0005	0.0346
Chile	890926	890929	556	Santiago IPSA	-0.1245***	0.1251	0.0002	0.0473	0.0014	0.0313
Mexico	810603	870102	699	Mexico IPC (Bolsa)	-0.1527***	0.1395	0.0021	0.0739	0.0020	0.0622
B. Asian-Pacific Funds: Emerging Market										
China	920710	920710	411	Shanghai SE	-0.0848***	0.1536	-0.0011	0.0486	-0.0003	0.0391
India Growth	880812	880812	615	Bombay SE National	-0.0584***	0.1967	0.0006	0.0423	-0.0003	0.0504
Indonesia	900301	900302	534	Jakarta SE Composite	0.1800***	0.1861	-0.0029	0.0644	-0.0032	0.0563
Korea	840822	870102	699	KOSPI	0.2109***	0.2693	0.0000	0.0606	0.0011	0.0446
Malaysia	870508	870508	681	Kuala Lumpur SE	0.0271***	0.2001	-0.0012	0.0646	-0.0009	0.0450
First Philippines	891108	891110	550	Philippines SE Composite	-0.1782***	0.1031	-0.0021	0.0498	-0.0013	0.0363
Taiwan	861223	870129	695	Taiwan SE Weighted	0.0703***	0.2552	-0.0001	0.0835	0.0010	0.0750
Thai	880217	880219	640	Bangkok SET	0.1598***	0.2632	-0.0021	0.0596	-0.0016	0.0500
Turkish Investment	891205	891208	546	Data Stream Price Index	-0.0123	0.1926	0.0007	0.0643	0.0012	0.0688
C. Asian-Pacific Funds: Developed Market										
First Australia	851212	870102	699	All Ordinaries	-0.1611***	0.0760	-0.0003	0.0455	-0.0004	0.0273
Japan OTC Equity	900314	900316	532	Nikkei 225 Index	0.0354***	0.1384	-0.0007	0.0526	0.0002	0.0364
Singapore	900724	900802	512	Straits Times Index	-0.0390***	0.1069	-0.0010	0.0442	-0.0005	0.0293
D. European Funds: Developed Market										
Austria	890921	890922	557	WBK Index	-0.1373***	0.1268	-0.0004	0.0495	0.0004	0.0347
France Growth	900510	900511	524	CAC 40 Index	-0.1659***	0.0755	0.0000	0.0381	0.0008	0.0250
Germany	860718	870102	699	DAX 100 Index	-0.0572***	0.1356	0.0004	0.0497	0.0004	0.0317
Italy	860226	870102	699	Milan MIB Index	-0.1278***	0.1065	0.0008	0.0476	0.0005	0.0296
Portugal	891101	891110	550	Portugal BVL General	-0.1117***	0.1038	-0.0006	0.0461	0.0001	0.0272
Spain	880621	880624	622	Madrid SE General	-0.0381***	0.2001	0.0000	0.0495	0.0003	0.0288
Swiss Helvetia	870819	870828	665	Swiss Market	-0.1051***	0.0910	0.0012	0.0374	0.0015*	0.0233

Note: The ***, **, and * indicate significance of mean value at 1%, 5%, and 10% levels, respectively.

Table 2
Fund Premiums as a Function of Expected Volatility and Trading Volume

$$PR_{i,t} = \alpha + \gamma_m V_{m,t} + \gamma_m^* V_{m,t}^* + \lambda_m \sigma_{m,t+1}^2 + \lambda_m^* \sigma_{m,t+1;LC}^2 + \delta_s \sigma_{s,t+1}^2$$

where $PR_{i,t}$ = weekly fund premium (log difference between fund price and NAV in U.S. dollars); $V_{m,t}$ = trading volume relative to market capitalization in the U.S. stock market; $V_{m,t}^*$ = trading volume relative to market capitalization in the foreign stock market; $\sigma_{m,t+1}^2$ = conditional variance of U.S. market return for time t+1 expected at time t; $\sigma_{m,t+1;LC}^2$ = conditional variance of foreign market return in local currency for time t+1 expected at time t; and $\sigma_{s,t+1}^2$ = conditional variance of foreign exchange return for time t+1 expected at time t.

Fund	Constant	$V_{m,t}$	$V_{m,t}^*$	$\sigma_{m,t+1}^2$	$\sigma_{m,t+1;LC}^2$	$\sigma_{s,t+1}^2$	R^2
A. Latin-American Funds: Emerging Market							
Argentina	0.0080 (0.29)	-0.0331*** (5.85)	0.0202 (1.35)	-347.7342*** (5.87)	4.6038 (0.24)	-96.4891 (0.11)	0.67
Chile	-0.0484 (1.57)	-0.0226*** (4.02)	-0.0115 (0.88)	-237.4968*** (6.51)	-13.1755 (0.48)	-344.2373 (1.04)	0.57
Mexico	-0.0330 (1.25)	-0.0420*** (8.88)	0.0068*** (2.94)	-480.4428*** (7.25)	62.5299*** (4.44)	3.5859*** (9.27)	0.73
B. Asian-Pacific Funds: Emerging Market							
China	-0.0500 (1.49)	-0.0188*** (4.56)	-0.0031*** (2.85)	-62.0965** (2.26)	0.3072*** (4.09)	3.5917 (0.49)	0.31
India Growth	-0.0414 (1.32)	-0.0214*** (2.95)	0.0325*** (4.50)	-337.3062*** (8.59)	8.3560 (0.51)	-4.1815 (0.26)	0.57
Indonesia	0.2683*** (4.28)	-0.0115 (1.31)	-0.0050 (0.42)	-19.3384*** (3.41)	-0.1107 (0.14)	4.7531*** (5.79)	0.21
Korea	0.0747* (1.88)	-0.0196* (1.83)	-0.0043** (2.46)	-193.5586*** (5.24)	28.1385*** (7.82)	5.0641*** (2.63)	0.62
Malaysia	0.1740* (1.87)	-0.0154 (0.85)	0.0087 (1.11)	86.2636** (2.09)	5.6283** (2.22)	-1.9753* (1.89)	0.16
First Philippines	-0.1810*** (8.80)	-0.0031 (0.66)	0.0101* (1.67)	-72.8288*** (2.80)	4.9895*** (4.52)	93.9558*** (6.68)	0.53
Taiwan	-0.0740 (1.25)	0.0118 (1.45)	-0.0057** (2.13)	-159.9517*** (3.94)	110.9889*** (4.80)	-234.2145* (1.91)	0.02
Thai	0.2323* (1.94)	0.0028 (0.13)	0.0038 (0.37)	-7.1386 (0.26)	-0.3070 (0.04)	25.6902* (1.79)	0.07
Turkish Investment	-0.0356 (1.05)	-0.0053 (0.99)	-0.0065*** (2.64)	-138.5016*** (2.81)	1.8605 (0.85)	36.5089 (1.21)	0.34
C. Asian-Pacific Funds: Developed Market							
First Australia	0.1167** (2.34)	-0.0157*** (4.94)	-0.0042 (1.60)	22.5781 (0.99)	-8.7994 (0.69)	-1185.7398*** (4.78)	0.25
Japan OTC Equity	0.0386 (0.77)	-0.0184 (1.26)	-0.0007*** (4.10)	-147.7200*** (3.78)	36.8944 (1.09)	159.2376*** (2.72)	0.23
Singapore	-0.0109 (0.28)	-0.0097 (0.94)	-0.0146 (1.29)	-135.0412*** (4.21)	-11.9272 (1.13)	238.1629*** (3.30)	0.19

Fund	Constant	$V_{m,t}$	$V_{m,t}^*$	$\sigma_{m,t+1}^2$	$\sigma_{m,t+1;LC}^{2*}$	$\sigma_{s,t+1}^2$	R^2
D. European Funds: Developed Market							
Austria	-0.0392 (0.98)	-0.0286*** (5.02)	0.0063** (1.98)	-7.1520 (0.18)	3.5727 (0.65)	-144.0465 (1.19)	0.19
France Growth	-0.1237*** (5.42)	-0.0096** (2.24)	0.0004 (0.09)	12.6869 (0.65)	-16.5978 (1.07)	-41.6995 (0.75)	0.03
Germany	0.0435 (1.08)	-0.0401*** (5.69)	-0.0003 (1.43)	264.1745*** (5.58)	-32.6481 (1.58)	140.0234 (1.57)	0.32
Italy	-0.0573** (1.96)	-0.0201*** (3.45)	0.0005 (0.20)	-13.7644 (0.96)	-36.0831** (2.01)	8.5520 (1.10)	0.17
Portugal	-0.0352 (1.02)	-0.0194*** (3.69)	0.0001 (0.14)	-9.7406 (0.27)	-19.4280*** (2.76)	-90.4829 (1.16)	0.11
Spain	0.1226*** (2.61)	-0.0489*** (5.66)	-0.0011 (0.31)	220.2651*** (3.66)	-26.5729 (1.47)	-28.6780 (1.34)	0.21
Swiss Helvetia	0.0512 (1.64)	-0.0380*** (8.96)	-0.0031 (1.18)	35.8615*** (3.18)	-50.2023*** (4.54)	-70.1187 (0.73)	0.46

Note: The ***, **, and * indicate significance of mean value at 1%, 5%, and 10% levels, respectively.

Table 3
Tests of Joint Significance for Fund Premiums

Tests are based on estimations in Table 2. In Panel A, for each variable, the first column contains chi-square statistics for testing the null of the same coefficients on U.S. and foreign market trading volumes, and the second column includes p-values associated with corresponding chi-square statistics. In Panel B, the numbers are chi-square statistics for testing the null of the same coefficients on the corresponding U.S. or foreign variable across funds in the same region, and numbers in parentheses are p-values associated with corresponding chi-square statistics. Emerging funds include Latin-American funds and emerging Asian-Pacific funds, and developed funds include developed Asian-Pacific funds and European funds.

Panel A: Test for the Same Coefficients on the U.S. and Foreign Variables for Each Fund

Fund	Trading Volume ($H_0: \gamma_m = \gamma_m^*$)		Market Volatility ($H_0: \lambda_m = \lambda_m^*$)	
	χ^2 statistics	p-value	χ^2 statistics	p-value
A. Latin-American Funds: Emerging Market				
Argentina	10.44***	0.001	26.71***	0.000
Chile	0.77	0.377	14.79***	0.000
Mexico	126.36***	0.000	48.40***	0.000
B. Asian-Pacific Funds: Emerging Market				
China	14.93***	0.000	5.12**	0.024
India Growth	26.79***	0.000	60.24***	0.000
Indonesia	0.16	0.687	10.11***	0.001
Korea	2.10	0.147	39.8***	0.000
Malaysia	1.76	0.185	3.70*	0.054
First Philippines	2.63	0.104	8.28***	0.004
Taiwan	3.76*	0.053	29.84***	0.000
Thai	0.00	0.968	0.07	0.792
Turkish Investment	0.03	0.852	8.43***	0.004
C. Asian-Pacific Funds: Developed Market				
First Australia	7.40***	0.007	2.33	0.127
Japan OTC Equity	1.49	0.221	14.15***	0.000
Singapore	0.07	0.783	14.53***	0.000
D. European Funds: Developed Market				
Austria	23.20***	0.000	0.06	0.801
France Growth	2.42	0.119	1.13	0.287
Germany	32.76***	0.000	24.34***	0.000
Italy	10.16***	0.001	2.44	0.118
Portugal	16.53***	0.000	0.07	0.790
Spain	35.25***	0.000	11.59***	0.000
Swiss Helvetia	45.29***	0.000	21.25***	0.000

Panel B: Test for the Same Coefficients on the U.S. and Foreign Variables across Funds in the Region

Region	US Trading Volume (γ_m 's)	Foreign Trading Volume (γ_m^* 's)	US Market Volatility (λ_m 's)	Foreign Market Volatility (λ_m^* 's)	Exchange Rate Volatility (δ_s 's)
Latin American	17.02*** (0.000)	3.86 (0.145)	10.36*** (0.006)	35.84*** (0.000)	1.22 (0.541)
Emerging Asian-Pacific	71.94*** (0.000)	38.18*** (0.000)	148.37*** (0.000)	115.91*** (0.000)	106.60*** (0.000)
Developed Asian-Pacific	0.86 (0.649)	3.09 (0.213)	17.63*** (0.000)	1.99 (0.369)	33.39*** (0.000)
European	85.57*** (0.000)	5.93 (0.431)	34.91*** (0.000)	48.05*** (0.000)	16.19** (0.013)
Emerging	54.36*** (0.000)	100.84*** (0.000)	266.79*** (0.000)	306.63*** (0.000)	178.01*** (0.000)
Developed	96.32*** (0.000)	17.76** (0.038)	56.73*** (0.000)	76.07*** (0.000)	45.19*** (0.000)
All	157.16*** (0.000)	231.39*** (0.000)	706.50*** (0.000)	712.19*** (0.000)	321.83*** (0.000)

Note: The ***, **, and * indicate significance of mean value at 1%, 5%, and 10% levels, respectively.

Table 4
Contagion of Emerging-Market Crises

$$PR_{i,t} = \alpha + \nu MEXICO + \omega ASIA + \gamma_m V_{m,t} + \gamma_m^* V_{m,t}^* + \lambda_m \sigma_{m,t+1}^2 + \lambda_m^* \sigma_{m,t+1;LC}^{2*} + \delta_s \sigma_{s,t+1}^2$$

where $PR_{i,t}$ = weekly fund premium (log difference between fund price and NAV in the U.S. dollar); $ASIA$ = dummy variable for 1994 Mexican peso crisis; value of 1 for the crisis period [1994:12:2 - 1995:12:29] and 0 otherwise; $MEXICO$ = dummy variable for 1997 Asian financial crisis; value of 1 for the crisis period [1997:7:4 - 2000:5:26] and 0 otherwise; $V_{m,t}$ = trading volume relative to market capitalization in the U.S. stock market; $V_{m,t}^*$ = trading volume relative to market capitalization in the foreign stock market; $\sigma_{m,t+1}^2$ = conditional variance of U.S. market return for time $t+1$ expected at time t ; $\sigma_{m,t+1;LC}^{2*}$ = conditional variance of foreign market return in the local currency for time $t+1$ expected at time t ; and $\sigma_{s,t+1}^2$ = conditional variance of foreign exchange return for time $t+1$ expected at time t . In Panel A, the numbers are estimated regression coefficients on the corresponding emerging-market crisis dummy variable, and numbers in the parentheses are t-statistics for testing the significance of the regression coefficients. In Panel B, the numbers are chi-square statistics for testing the null of the same coefficients on the corresponding emerging-market crisis dummy variable across funds in the same region, and numbers in parentheses are p-values associated with corresponding chi-square statistics. Emerging funds include Latin-American funds and emerging Asian-Pacific funds, and developed funds include developed Asian-Pacific funds and European funds.

Panel A: Regression Coefficients on Dummy Variables for Emerging Market Crises

Fund	MEXICO	ASIA
A. Latin-American Funds: Emerging Market		
Argentina	0.0298 (1.17)	-0.1800*** (5.71)
Chile	-0.0834*** (4.24)	-0.1402*** (6.32)
Mexico	0.0763** (2.16)	-0.1344*** (6.52)
B. Asian-Pacific Funds: Emerging Market		
China	0.1299*** (2.81)	-0.1124*** (2.91)
India Growth	-0.0168 (0.46)	-0.2260*** (5.90)
Indonesia	0.1624*** (2.85)	0.3018*** (4.33)
Korea	-0.0793*** (6.82)	-0.1086*** (3.69)
Malaysia	0.0197 (0.86)	0.3981*** (7.86)
First Philippines	-0.0028 (0.23)	0.0616*** (3.84)
Taiwan	0.0105 (0.18)	-0.2494*** (4.71)
Thai	-0.1796*** (4.15)	0.5368*** (7.54)
Turkish Investment	0.0444 (0.84)	-0.1618*** (5.24)

Fund	MEXICO	ASIA
C. Asian-Pacific Funds: Developed Market		
First Australia	-0.0356** (2.05)	-0.0695*** (4.33)
Japan OTC Equity	0.0046 (0.16)	0.0351 (0.89)
Singapore	0.0641** (2.51)	-0.0420 (1.07)
D. European Funds: Developed Market		
Austria	-0.0702*** (3.34)	-0.0123 (0.53)
France Growth	-0.0378** (2.02)	-0.0353* (1.79)
Germany	-0.1042*** (4.45)	-0.1090*** (4.11)
Italy	-0.0524** (2.45)	-0.0810*** (3.57)
Portugal	-0.0144 (0.66)	-0.0473 (1.58)
Spain	-0.1045*** (4.00)	-0.0884** (2.40)
Swiss Helvetia	-0.0198 (1.21)	-0.1125*** (5.75)

Panel B: Joint Significance Tests for the Same Coefficients on Crisis Dummy Variables across the Funds in the Region

Region	MEXICO (ν 's)	ASIA (ω 's)
Latin American	21.93*** (0.000)	7.03** (0.030)
Emerging Asian-Pacific	103.84*** (0.000)	118.48*** (0.000)
Developed Asian-Pacific	29.07*** (0.000)	7.27** (0.026)
European	44.58*** (0.000)	49.93*** (0.000)
Emerging	406.05*** (0.000)	213.68*** (0.000)
Developed	56.47*** (0.000)	70.43*** (0.000)
All	1120.07*** (0.000)	547.26*** (0.000)

Note: The ***, **, and * indicate significance of mean value at 1%, 5%, and 10% levels, respectively.

Table 5
Estimation Results for Fund Premiums: An Augmented Model

$$PR_{i,t} = \alpha + \gamma_m V_{m,t} + \gamma_m^* V_{m,t}^* + \lambda_m \sigma_{m,t+1}^2 + \lambda_m^* \sigma_{m,t+1;LC}^{2*} + \delta_s \sigma_{s,t+1}^2 + \beta_{1m}^* R_{m,t-1}^* + \beta_{2m}^* R_{m,t}^* + \beta_{3m}^* R_{m,t+1}^* + \phi_{1s} \Delta S_{t-1} + \phi_{2s} \Delta S_t + \phi_{3s} \Delta S_{t+1} + \beta_m R_{m,t}$$

where $PR_{i,t}$ = weekly fund premium (log difference between fund price and NAV in U.S. dollars); $V_{m,t}$ = trading volume relative to market capitalization in the U.S. stock market; $V_{m,t}^*$ = trading volume relative to market capitalization in the foreign stock market; $\sigma_{m,t+1}^2$ = conditional variance of U.S. market return for time $t+1$ expected at time t ; $\sigma_{m,t+1;LC}^{2*}$ = conditional variance of foreign market return in local currency for time $t+1$ expected at time t ; $\sigma_{s,t+1}^2$ = conditional variance of foreign exchange return for time $t+1$ expected at time t ; $R_{m,t}^*$ = foreign stock market returns in the local currency at time t ; ΔS_t = foreign exchange returns at time t ; and $R_{m,t}$ = US stock market returns in U.S. dollars at time t .

Fund	Constant	$V_{m,t}$	$V_{m,t}^*$	$\sigma_{m,t+1}^2$	$\sigma_{m,t+1;LC}^{2*}$	$\sigma_{s,t+1}^2$	$R_{m,t-1}^*$	$R_{m,t}^*$	$R_{m,t+1}^*$	ΔS_{t-1}	ΔS_t	ΔS_{t+1}	$R_{m,t}$	R^2
A. Latin-American Funds: Emerging Market														
Argentina	0.0067 (0.24)	-0.0331*** (5.97)	0.0223 (1.44)	-349.19*** (5.85)	4.8823 (0.25)	-267.60 (0.27)	0.0385 (0.61)	0.0160 (0.17)	-0.0387 (0.58)	-0.4401 (0.08)	-1.4218 (0.27)	1.5993 (0.34)	0.0472 (0.21)	0.67
Chile	-0.0485 (1.54)	-0.0221*** (3.92)	-0.0151 (1.10)	-247.46*** (7.14)	-10.534 (0.41)	-357.87 (1.20)	0.1171 (1.02)	0.1482 (1.14)	-0.0551 (0.38)	-0.0879 (0.20)	-0.5159 (0.93)	-0.5046 (0.99)	-0.1367 (0.72)	0.57
Mexico	-0.0324 (1.26)	-0.0423*** (9.27)	0.0068 (3.02)	-477.05*** (7.07)	62.037*** (4.22)	3.2097*** (6.13)	0.1576* (1.83)	0.0758 (0.86)	-0.0424 (0.47)	-0.1507 (1.57)	-0.1171 (0.60)	-0.1580* (1.76)	-0.2190 (1.11)	0.73
B. Asian-Pacific Funds: Emerging Market														
China	-0.0486 (1.53)	-0.0182*** (4.64)	-0.0048*** (3.89)	-59.872** (2.11)	0.2938*** (3.90)	11.188 (1.25)	-0.0102 (0.15)	0.1385* (1.76)	0.0349 (0.42)	-3.4275 (0.85)	8.0183** (2.34)	5.5231 (1.30)	0.1567 (0.78)	0.31
India Growth	-0.0605** (2.04)	-0.0181*** (2.64)	0.0298*** (4.33)	-303.25*** (7.61)	-2.7710 (0.19)	12.948 (0.80)	0.4321*** (4.24)	-0.0043 (0.05)	-0.0900 (1.32)	-0.3196 (0.73)	-0.6197 (1.28)	-0.1792 (0.53)	0.3557** (2.35)	0.55
Indonesia	0.2584*** (4.32)	-0.0115 (1.28)	0.0011 (0.10)	-34.089** (2.54)	-0.2009 (0.26)	5.2620*** (6.28)	0.1764* (1.72)	-0.1531 (1.07)	-0.1665 (0.82)	0.0755 (0.84)	-0.4057 (1.56)	0.0799 (0.32)	-0.1572 (0.54)	0.24
Korea	0.0733** (2.04)	-0.0194** (2.05)	-0.0042** (2.54)	-173.66*** (4.99)	33.991*** (8.34)	0.9009 (0.36)	0.3276*** (3.55)	0.3005*** (2.72)	0.2560*** (2.99)	-0.4866*** (4.62)	-0.4990*** (2.90)	-0.3592*** (2.74)	0.2293 (1.39)	0.65
Malaysia	0.1913** (2.10)	-0.0187 (1.08)	0.0110 (1.58)	75.001* (1.79)	9.8310*** (3.05)	-3.6723*** (2.68)	0.6515*** (3.43)	0.0854 (0.42)	0.3861* (1.92)	-0.7644*** (2.98)	-0.5455* (1.92)	-0.3185 (1.09)	0.0809 (0.21)	0.15
First Philippines	-0.1826*** (8.56)	-0.0027 (0.55)	0.0096 (1.55)	-62.548*** (2.63)	4.7645*** (4.64)	83.968*** (6.80)	0.1361** (2.22)	-0.0653 (0.67)	-0.0412 (0.45)	-0.2643 (1.34)	-0.2140 (0.76)	-0.1160 (0.52)	0.1636 (1.02)	0.54
Taiwan	-0.0521 (0.94)	0.0124* (1.66)	-0.0087*** (3.16)	-152.27*** (4.11)	110.43*** (4.95)	-247.47** (2.13)	0.4472** (2.57)	0.6909*** (2.97)	0.4168** (2.57)	0.5928 (0.92)	0.0815 (0.11)	0.3497 (0.61)	0.1357 (0.74)	0.14
Thai	0.2580** (2.25)	-0.0027 (0.13)	0.0051 (0.69)	-0.5928 (0.02)	2.4443 (0.29)	13.365 (0.80)	0.2643 (1.54)	-0.0197 (0.11)	0.0757 (0.48)	0.3877 (1.23)	-0.4109 (1.22)	-0.2495 (0.70)	0.3410 (0.79)	0.06
Turkish Investment	-0.0618 (1.64)	-0.0039 (0.73)	-0.0066 (2.65)	-158.80*** (3.31)	1.7813 (0.79)	40.923 (1.36)	0.1680*** (3.21)	0.0057 (0.11)	-0.0050 (0.14)	-0.5767 (1.12)	-0.8224 (1.54)	-0.2509 (0.50)	0.2327 (1.14)	0.39

Fund	Constant	$V_{m,t}$	$V_{m,t}^*$	$\sigma_{m,t+1}^2$	$\sigma_{m,t+1;LC}^{2*}$	$\sigma_{s,t+1}^2$	$R_{m,t-1}^*$	$R_{m,t}^*$	$R_{m,t+1}^*$	ΔS_{t-1}	ΔS_t	ΔS_{t+1}	$R_{m,t}$	R^2
C. Asian-Pacific Funds: Developed Market														
First Australia	0.1207*** (2.59)	-0.0158*** (5.13)	-0.0045 (1.64)	20.386 (0.90)	-6.4298 (0.50)	-1207.40*** (5.23)	0.2379 (1.45)	0.0351 (0.18)	0.1158 (0.85)	-0.1582 (0.70)	0.3091 (1.32)	0.1303 (0.67)	0.0861 (0.73)	0.26
Japan OTC Equity	0.0465 (0.93)	-0.0194 (1.33)	-0.0008*** (4.54)	-131.35*** (3.90)	30.442 (0.92)	153.43** (2.42)	0.2800* (1.70)	0.6870*** (4.47)	0.3268* (1.90)	0.1828 (0.62)	-0.5868** (2.02)	-0.3713 (1.37)	0.3129 (1.29)	0.27
Singapore	-0.0145 (0.37)	-0.0093 (0.89)	-0.0129 (1.12)	-138.74*** (4.54)	-13.639 (1.40)	228.92*** (3.22)	0.1430 (1.13)	-0.0617 (0.46)	-0.0771 (0.72)	0.0820 (0.18)	-0.1804 (0.39)	0.1651 (0.32)	0.3423* (1.92)	0.20
D. European Funds: Developed Market														
Austria	-0.0371 (0.98)	-0.0289*** (5.20)	0.0068** (2.19)	-3.8516 (0.10)	2.0170 (0.38)	-154.69 (1.31)	0.2090* (1.70)	-0.2520** (2.38)	-0.0812 (0.90)	0.4869*** (3.05)	-0.2574 (1.48)	0.0330 (0.17)	0.0463 (0.27)	0.19
France Growth	-0.1223*** (5.42)	-0.0099** (2.36)	0.0006 (0.15)	8.6743 (0.44)	-12.718 (0.82)	-49.318 (0.85)	0.0353 (0.45)	-0.0987 (1.00)	0.0428 (0.52)	0.2317 (1.17)	-0.3689* (1.74)	-0.0009 (0.01)	0.1494 (0.84)	0.03
Germany	0.0442 (1.13)	-0.0402*** (5.74)	-0.0002 (1.00)	265.39*** (5.72)	-33.935 (1.63)	140.13 (1.57)	0.0646 (0.69)	-0.2617** (2.07)	-0.1074 (1.10)	0.4605** (2.53)	-0.1616 (0.88)	-0.1369 (0.65)	-0.0943 (0.55)	0.32
Italy	-0.0541* (1.87)	-0.0208*** (3.63)	0.0013 (0.54)	-17.891 (1.28)	-35.229** (1.96)	-8.5338 (0.85)	0.1064 (1.21)	-0.1019 (1.17)	-0.0486 (0.65)	0.1556 (0.77)	-0.4962** (2.08)	-0.3917* (1.93)	-0.0352 (0.18)	0.17
Portugal	-0.0317 (0.93)	-0.0197*** (3.81)	0.0004 (0.41)	-10.447 (0.29)	-19.561*** (2.71)	-102.82 (1.34)	0.0709 (0.64)	-0.0000 (0.00)	-0.0443 (0.56)	0.3947** (1.96)	-0.3066 (1.41)	-0.0609 (0.33)	-0.0575 (0.40)	0.11
Spain	0.1278*** (2.66)	-0.0506*** (5.62)	0.0008 (0.24)	224.95*** (3.67)	-29.159 (1.55)	-36.815 (1.41)	0.1062 (0.93)	-0.3240*** (2.81)	-0.1719 (1.48)	0.4145* (1.77)	-0.1521 (0.47)	-0.0824 (0.26)	-0.1150 (0.62)	0.21
Swiss Helvetia	0.0614* (1.77)	-0.0379*** (9.38)	-0.0034 (1.28)	36.522*** (3.07)	-49.526*** (4.57)	-104.17 (0.97)	0.0222 (0.32)	-0.2260** (2.50)	-0.0603 (0.97)	0.3501*** (2.81)	-0.3535** (2.47)	0.0584 (0.44)	0.0247 (0.14)	0.47

Note: The ***, **, and * indicate significance of mean value at 1%, 5%, and 10% levels, respectively.

Table 6
Summary of Estimation Results of the Augmented Model

$$PR_{i,t} \text{ (or } R_{i,t} \text{ or } R_{i,t}^*) = \alpha + \gamma_m V_{m,t} + \gamma_m^* V_{m,t}^* + \lambda_m \sigma_{m,t+1}^2 + \lambda_m^* \sigma_{m,t+1;LC}^{2*} + \delta_s \sigma_{s,t+1}^2 + \beta_{1m}^* R_{m,t-1}^* + \beta_{2m}^* R_{m,t}^* + \beta_{3m}^* R_{m,t+1}^* + \phi_{1s} \Delta S_{t-1} + \phi_{2s} \Delta S_t + \phi_{3s} \Delta S_{t+1} + \beta_m R_{m,t}$$

where $PR_{i,t}$ = weekly fund premium (log difference between fund price and NAV in U.S. dollars); $R_{i,t}$ = weekly fund returns; $R_{i,t}^*$ = weekly NAV returns; $V_{m,t}$ = trading volume relative to market capitalization in the U.S. stock market; $V_{m,t}^*$ = trading volume relative to market capitalization in the foreign stock market; $\sigma_{m,t+1}^2$ = conditional variance of U.S. market return for time $t+1$ expected at time t ; $\sigma_{m,t+1;LC}^{2*}$ = conditional variance of foreign market return in local currency for time $t+1$ expected at time t ; $\sigma_{s,t+1}^2$ = conditional variance of foreign exchange return for time $t+1$ expected at time t ; $R_{m,t}^*$ = foreign stock market returns in the local currency at time t ; ΔS_t = foreign exchange returns at time t ; and $R_{m,t}$ = US stock market returns in U.S. dollars at time t . The number under each specification is the number of significant coefficients at the 10% significance level, and the number in the parentheses is the number of positive significant coefficients.

Variables	Fund Premiums	Fund Returns	NAV Returns
Constant	11 (7)	6 (0)	7 (1)
$V_{m,t}$	15 (1)	3 (2)	8 (6)
$V_{m,t}^*$	6 (2)	5 (5)	5 (3)
$\sigma_{m,t+1}^2$	16 (4)	2 (2)	7 (3)
$\sigma_{m,t+1;LC}^{2*}$	9 (6)	5 (1)	9(3)
$\sigma_{s,t+1}^2$	8 (5)	5 (5)	9 (7)
$R_{m,t-1}^*$	10 (10)	7 (7)	6 (6)
$R_{m,t}^*$	8 (4)	21 (21)	21 (21)
$R_{m,t+1}^*$	4 (4)	8 (5)	5 (5)
ΔS_{t-1}	7 (5)	1 (0)	7 (4)
ΔS_t	7 (1)	15 (14)	19 (18)
ΔS_{t+1}	3 (0)	10 (9)	8 (4)
$R_{m,t}$	2 (2)	21 (21)	5 (5)