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The effect of country-specific index trading on closed-end country funds: an empirical analysis

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Abstract

The American Stock Exchange initiated trading in 17 World Equity Benchmark Shares (acronym "WEBS"™) in April 1996. WEBS are index funds designed to track the Morgan Stanley Capital International (MSCI) indexes. We examine the effect of this event on closed-end country funds (CECFs) and find that percentage discounts increase. CECFs with a corresponding WEBS index experience the largest increase and also show a decline in trading volume. We attribute these results to (1) the effects of increased competition and (2) a reduction in the market segmentation premium. Since additional WEBS have begun trading, and other approvals may follow, similar effects could be experienced in the future. © 2001 Elsevier Science Inc. All rights reserved.

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1. Introduction

Closed-end funds are mutual funds that issue a fixed number of shares that trade on a stock exchange. Owners of closed-end funds liquidate their shares by selling them to other investors. Whereas open-end funds stand ready to redeem shares at net asset value (NAV), closed-end funds usually sell at a discount and sometimes at a premium to NAV. However, when the funds are initiated, they always sell at a premium. The start-up premium results

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from underwriting fees and up-front costs associated with the flotation. Attempts to fully explain discounts or premiums on a rational basis have generally failed, and this has become known as the "closed-end fund puzzle."

Nevertheless, it is not surprising that closed-ending is often more attractive than open-ending for funds specializing in a foreign country, i.e. funds in which the underlying assets trade in a foreign market. Such closed-end country funds (CECFs) may be illiquid, and closed-ending ensures that assets will not have to be sold off at fire sale prices to meet redemptions during market declines. Hence, CECFs represent rational international investment vehicles for many domestic investors.

Beginning April 5, 1996, the American Stock Exchange initiated trading in World Equity Benchmark shares (acronym WEBS™) designed to track the Morgan Stanley Capital International (MSCI) indexes for 17 countries. MSCI indexes have been the traditional benchmarks for international portfolios. Each WEBS index series holds a representative sample of the underlying securities in a corresponding MSCI index. Since WEBS can be sold to the fund or redeemed by the fund in aggregations of securities called "creation units," WEBS can be easily arbitraged and therefore track closely the values of their respective MSCI index. However, irregularities in portfolio sampling techniques and other regulatory constraints (mainly tax-related) allow WEBS values to vary slightly from MSCI indexes.

Few open-end mutual funds specialize in a single foreign market. Hence, prior to the arrival of international index funds, CECFs were often the only economical avenue for individual investors to diversify into single market foreign investments. Therefore, WEBS are an important integrating event and should impact CECF discounts. It should be noted that the week prior to the introduction of WEBS, *Deutsche Morgan Grenfell* introduced eight foreign Country Basket Index funds (CBs). Although quite similar to WEBS, investors never fully accepted CBs, and the series was discontinued on February 10, 1997, less than a year after trading began. However, as an integrating product, CBs may have had a similar impact on CECF discounts. Nevertheless, since CBs were introduced at almost exactly the same time as WEBS, were never accepted by investors, and have similar impact, we couch our discussion in terms of WEBS.

This paper uses a modified event study methodology to examine the impact of WEBS trading on mean CECF percentage discounts. We divide CECFs into two groups: those with and those without a corresponding (same country) WEBS index. We separately estimate the impact of WEBS trading on each group. We construct tests to examine the short-, intermediate-, and long-run impact of WEBS trading. We also use a two-factor difference in means test to examine the differential impact on percentage discounts for CECFs with and without a corresponding WEBS Index. Finally, we use the same methodologies to examine the impact of WEBS on trading volumes.

2. Theory and literature review

Considerable effort has been spent to explain deviations between closed-end fund prices and NAVs. Explanations in the literature fall into two categories: (1) behavioral theories, which focus on the irrational expectations of individual investors, and (2) economic theories that attempt to explain discounts within the framework of the efficient market hypothesis.

Among the behavioral proponents, Delong et al. (1990) and Lee, Shleifer, and Thaler (1991) attribute fluctuations of the discount or premium to changes in "investor sentiment." These authors suggest that the closed-end fund market in the U.S. is characterized by negligible institutional participation and is dominated by individual investors, who are less likely to trade on fundamentals. The behavioral model suggests two possible reactions to the introduction of WEBS trading. Investors may fail to appreciate the importance of WEBS trading, in which case no impact on discounts will be found. Alternatively, investors may irrationally bid down CECF values.

In contrast, economic explanations suggest rational premiums and discounts. The four classic economic explanations include: (1) biases in NAVs, (2) tax timing issues, (3) agency costs, and (4) international market segmentation. NAV biases reflect differences between personal and fund accounting of capital gains liabilities. They also reflect rational discounting of some portfolio assets due to liquidity concerns. However, Neal and Wheatley (1998) find no evidence of a liquidity effect in closed-end fund discounts. The tax timing explanation states that closed-end fund investors give up valuable tax-timing advantages. Seyhun and Skinner (1994) and Brickley, Manaster, and Schalheim (1991) find empirical support for a tax-timing effect. The agency cost explanation suggests that discounts are, in part, a rational response to excessive management fees, poor management performance, and managerial entrenchment. Pontiff (1995) and Barclay, Holderness, and Pontiff (1993) find evidence that agency costs are reflected in closed-end fund prices. Richard and Wiggins (2000) find that CECF premiums are able to predict NAV returns after controlling for the return on foreign markets, indicating that at least part of the discount reflects rational assessments of managerial abilities. Finally, the market segmentation explanation suggests that if a closed-end fund provides access to a restricted market, fund prices may reflect a rational premium. The restrictions need not be governmentally imposed. They may simply reflect the cost to domestic investors of creating well-diversified, single-country foreign investments. Bosner-Neal et al. (1990), Errunza (1991), and Levy-Yeyati and Ubide (1998) find evidence of a market segmentation premium in CECF prices.

Economic explanations suggest WEBS trading should tend to reduce market segmentation premiums, and the effect should be more evident in funds with a corresponding WEBS index. Further, some investors may abandon excessive fee CECFs for lower fee WEBS, putting downward pressure on prices and increasing discounts. Again, the effect would tend to be more evident in CECFs with a corresponding WEBS index. Unlike an overreaction, these changes would tend to be permanent.

The primary motivation for our paper is to empirically examine the effect of WEBS trading on CECF discounts. Our tests also shed light on theoretical explanations for closed-end fund discounts. In particular, the behavioral model will be supported by either: (1) no change in CECF discounts, or (2) a reversal pattern in which discounts initially increase but subsequently fall. On the other hand, a permanent increase in discounts favors the economic model. Since our tests on percentage discounts do not specifically differentiate between a market segmentation premium effect and the agency cost effect, we also examine the impact on CECF trading volume. If the agency explanation is correct, an impact on CECF volume should be evident.

3. Data and methodology

The data for this study are from the *Wiesenberger Investment Company*. Our sample comprises weekly closing market prices and NAVs. The data also include average daily trading volumes (in shares transacted) for each week in the study. The discount is determined as the NAV minus the price of the fund, and the percentage discount is defined as the discount divided by the NAV. The data are from all CECFs listed on the New York Stock Exchange and American Stock Exchange. We also include a country-specific Canadian fund that trades on the Toronto Stock Exchange, because a Canadian WEBS is listed. All of the funds are categorized by the *Wall Street Journal* (WSJ) as "Specialized Equity Funds." We limit this study to country specific funds. We exclude blended funds (funds that invest in more than one country) because the impact of WEBS trading may be ambiguous. To be included in the study we require funds to trade over the 160-week time period from March 18, 1994 through April 4, 1997.

Since the introduction of a WEBS index may have a stronger impact on a CECF from the same country, we divide the sample into funds with and funds without a corresponding WEBS index. For CECFs trading over the 160-week period, we find 18 funds with and 25 funds without a corresponding WEBS index. Appendix A lists the CECFs by group and shows the corresponding WEBS index for the first group. Table 1 presents mean percentage discounts and grand average daily volumes for the 160-week period. Note that volume data were unavailable for 4 funds.

3.1. Modified event study

The traditional approach to measuring event impact is cumulative abnormal residual (CAR) analysis. However, Campbell, Lo, and Mackinlay, (1997 henceforth CLM) caution against using CAR analysis when events are clustered. Event clustering results in cross-correlations among residuals, violating the normal CAR assumption of independence. Since we examine the impact of a simultaneous event on all funds, we clearly have clustering. CLM suggest a system of equations model used by Shipper and Thompson (1983 henceforth ST), which we use below.

The pre-event estimation window includes the 104-week period from March 18, 1994 through March 8, 1996. The four weeks culminating with the April 5, 1996 introduction of WEBS trading are the event window. Observations during the event window are excluded to avoid unintentional biases in the estimation of post-event mean discounts. The remaining 52 weeks from April 12, 1996 through April 4, 1997 comprise the post-event evaluation period. We subdivide the post-event evaluation period into overlapping 13-, 26-, and 52-week periods.

The model is constructed as a system of equations with each individual equation corresponding to a particular fund. We define T as the number of observations per equation. In particular, T will equal the sum of pre- and post-event weeks. For example, if the post-event window is 13 weeks, then T will equal $52 + 13 = 65$ observations. Let Y_j be a $T \times 1$ vector of observed percentage discounts for firm j. Let D be a $T \times 1$ vector of dummy variables with zeroes in the pre-event estimation period and ones in the post-event test period. Finally, let ε_j be a $T \times 1$ vector of serially independent error terms for firm j. Then for each individual firm, the model can be written as:

$$Y_j = \mu_j + \beta_j D + \varepsilon_j \text{ with } \varepsilon_j \sim N(0, \sigma^2) \tag{1}$$

Let J equal the number of firms in a particular group. Then, following ST, the J individual equations are stacked into a system of equations as:

$$\mathbf{Y} = \mathbf{X}\Gamma + \mathbf{E} \tag{2}$$

where:

$$Y = \begin{bmatrix} Y_1 \\ \vdots \\ Y_J \end{bmatrix}, X = \begin{bmatrix} \bar{X} & 0 \\ 0 & \bar{X} \end{bmatrix}, \bar{X} = [1 \quad D], E = \begin{bmatrix} \varepsilon_1 \\ \vdots \\ \varepsilon_j \end{bmatrix}, \text{and } \Gamma = \begin{bmatrix} \alpha_1 \\ \beta_1 \\ \vdots \\ \alpha_J \\ \beta_J \end{bmatrix}$$

The model is robust to correlation of the ε_j across the J equations. Hence, it is appropriate for clustered events. With a known residual covariance matrix Σ, GLS provides the best unbiased maximum likelihood estimate of Γ. Using an OLS estimate for Σ, the joint GLS estimate of Γ is consistent and asymptotically efficient. Since the explanatory variables are identical across equations, OLS provides identical parameter estimates as GLS.

For all tests, the null hypothesis is that the introduction of WEBS has no effect on percentage discounts. A positive and significant β_j coefficient indicates a fund's mean percentage discount increases after the introduction of WEBS. However, following ST, we also construct two sets of hypotheses regarding the joint impact of WEBS trading. The first joint hypothesis is that the sum of the β_j equal zero:

$$H_0: \sum_{j=1}^{J} \beta_j = 0 \tag{3}$$

Paraphrasing ST, this test is analogous to testing the sample-wide abnormal returns in a traditional CAR methodology. We also test the hypothesis that the individual β_j parameters are jointly zero:

$$H_0: \beta_j = 0 \; \forall j \tag{4}$$

This test may be useful if some funds react positively while others react negatively, hence canceling each other out in the sum. ST give the following test statistic to evaluate the two cross-equation hypotheses:

$$(a - A\hat{\Gamma})'[A(X'(\hat{\Sigma}^{-1} \otimes I)X)^{-1}A']^{-1}(a - A\hat{\Gamma}) \tag{5}$$

In the first joint test $a = 0$, and A is a row vector of ones and zeros. In the second test, a is a column vector of zeros, and A is a matrix of ones and zeros. As T goes to infinity, (5) is distributed as $\chi^2(q)$ in the limit, where q represents the number of restrictions. Note that the hypothesis test from equation (3) has one restriction. The hypothesis test in equation (4) will have $q = J$ restrictions, because the system has J separate equations, one for each CECF.

We separately estimate the system of equations model for each group of funds (i.e., those with and without corresponding WEBS). Further, we estimate separate models for the three post-event time periods. In the first run, we use a 13-week post-event window. This indicates the short run impact of WEBS trading. To examine the persistence of fund discount responses over the intermediate and long run, we also estimate the models using 26- and 52-week post-event windows.

3.2. The two-factor difference in means test

Extending our analysis, we also conduct a two-factor difference in means test. This allows us to examine whether WEBS had a significantly larger impact on corresponding CECFs. The model's first factor is time and has two levels: the pre- and post-event observations. The second factor is fund group: CECFs with and without a corresponding WEBS index. To construct the model, again let Y be a fund's observed percentage discount. The dummy regression is:

$$Y = \mu + \alpha D_1 + \gamma D_2 + \Theta D_1 D_2 + \varepsilon \qquad \varepsilon \sim N(0, \sigma^2) \tag{6}$$

where:

$$D_1 = \begin{cases} 1 \text{ if the observation is in the pre-event window} \\ -1 \text{ if the observation is in the post-event window} \end{cases}$$

$$D_2 = \begin{cases} 1 \text{ if the observation is from a CECF with a corresponding WEBS index} \\ -1 \text{ if the observation is from a CECF with no corresponding WEBS Index} \end{cases}$$

We estimate the model as a single regression on a pooled sample of percentage discounts from all 43 funds. Hence, μ is the average discount over the entire sample period for all funds. The α parameter measures the effect due to the time factor, γ measures the effect due to fund type, and Θ measures the interaction effect between the two factors. In order to investigate the differential impact on CECFs, let the subscript i equal 1 for a pre-event observation and 2 otherwise. Let the subscript k equal 1 for an observation from a CECF with a corresponding WEBS and 2 otherwise. Then the mean percentage discount (μ_{ik}) for each level of each factor can be recovered from the model via substitution. For example:

Table 1
Descriptive Statistics

Fund Name	Mean Percentage Discount	Mean Daily Volume
Funds with a Corresponding WEBS		
Austria Fund	16.9%	22,097
Canadian General Investments Ltd	23.7%	NA
Emerging Mexico Fund	2.2%	NA
First Australia Fund	14.1%	45,420
France Growth Fund	17.1%	28,544
Germany Fund	16.4%	44,572
Growth Fund of Spain	18.5%	NA
Italy Fund	12.2%	32,964
Japan Equity Fund	−8.1%	54,402
Japan OTC Equity Fund	−4.8%	47,528
Malaysia Fund	5.7%	31,929
Mexico Equity & Income Fund	2.2%	50,991
Mexico Fund	6.4%	317,097
New Germany Fund	21.0%	87,823
Singapore Fund	−1.7%	23,921
Spain Fund	15.0%	27,183
Swiss Helvetia Fund	10.5%	27,619
United Kingdom Fund	15.5%	NA
Funds without a Corresponding WEBS		
Argentina Fund	0.4%	49,137
Brazil Fund	4.4%	86,095
Brazilian Equity Fund	2.2%	45,249
Chile Fund	9.7%	39,204
China Fund	−1.8%	40,359
First Israel Fund	2.9%	40,211
Fidelity Philippine Fund	18.8%	130,587
Greater China Fund	6.9%	34,686
India Fund	6.5%	14,098
India Growth Fund	−7.0%	11,471
Irish Investment Fund	13.3%	34,384
Jardine Fleming China Region Fund	5.2%	124,198
Jardine Fleming India Fund	1.1%	31,099
Korea Equity Fund	3.4%	37,033
Korea Fund	−9.1%	136,821
Korean Investment Fund	3.0%	39,131
Morgan Stanley Indian Fund	1.9%	55,896
New South Africa Fund	18.8%	76,657
Pakistan Investment Fund	12.8%	39,120
ROC Taiwan Fund	0.0%	25,720
Taiwan Fund	−2.9%	19,338
Templeton China World Fund	7.3%	78,475
Thai Capital Fund	6.7%	48,554
Thai Fund	6.4%	27,581
Turkish Investment Fund	−15.6%	28,825

Calculations based on 160 total observations.
NA = not available.

$$\begin{cases}\mu_{11} = \mu - \alpha - \gamma + \Theta \\ \mu_{21} = \mu + \alpha - \gamma - \Theta \\ \mu_{12} = \mu - \alpha + \gamma - \Theta \\ \mu_{22} = \mu + \alpha + \gamma + \Theta\end{cases} \tag{7}$$

The relevant null hypotheses are:

1. H_0: $\mu_{21} - \mu_{11} = 0$ The change in the mean percentage discount for funds with a corresponding WEBS index is zero.
2. H_0: $\mu_{22} - \mu_{12} = 0$ The change in the mean percentage discount for funds with no corresponding WEBS index is zero.
3. H_0: $\mu_{21} - \mu_{11} - (\mu_{22} - \mu_{12}) = 0$ The change in the mean percentage discount for funds with a corresponding WEBS index is no different than for funds without a corresponding WEBS index.

Coefficient restriction tests of these hypotheses can be constructed from the definitions of the μ_{ik}.

4. Results

4.1. Results of the modified event study

Table 2—Panel A presents results for the J = 18 funds with corresponding WEBS indexes. For each individual equation, the intercept shows a fund's mean percentage discount over the pre-event estimation window. Each fund's intercept remains constant, regardless of the length of the post-event window. The β_j coefficients estimate the change in the mean percentage discount from the pre- to the post-event period. For the 13-week window, 15 of the 18 funds (83%) have positive β_j coefficients, with 14 of the positive estimates significant at the 1% level. The average change in percentage discount is 6.8%. More importantly, the hypothesis that the sum of the β_j coefficients equals zero is strongly rejected. This result is analogous to finding significant CARs in traditional event study methods. Not surprisingly, the hypothesis that the β_j coefficients are jointly zero is also strongly rejected.

The 26-week and 52-week post-event runs confirm that the increase in percentage discount persists. In the 26-week run, 17 of the 18 funds (94%) have positive β_j estimates. Of these, 15 are significant at the 1% level. The average change in percentage discount is 7.7%, and not surprisingly, both joint hypotheses of no change in percentage discount are strongly rejected. Similar results occur in the 52-week run.

The significant and persistent increase in percentage discounts is consistent with an economic explanation of closed-end fund discounts, and inconsistent with a behavioral

Table 2
Modified Event Study Results (Dependent Variable = Percentage Discount)

Fund Number	Fund Name	Intercept	$\hat{\beta}_j$ for post-event window of: 13-Weeks	26-Weeks	52-Weeks
Panel A: Funds With Corresponding WEBS					
1	Austria Fund	14.5%**	6.7%**	7.2%**	6.8%**
2	Canadian General Investments Ltd	25.4%**	−0.3%	−1.5%*	−5.1%**
3	Emerging Mexico Fund	−6.0%**	22.1%**	22.7%**	23.6%**
4	First Australia Fund	12.4%**	4.2%**	4.4%**	4.8%**
5	France Growth Fund	15.4%**	4.2%**	4.7%**	5.1%**
6	Germany Fund	15.0%**	3.7%**	4.6%**	4.0%**
7	Growth Fund of Spain	18.3%**	−0.7%	0.2%	0.7%
8	Italy Fund	9.9%**	5.3%**	6.3%**	6.7%**
9	Japan Equity Fund	−8.1%**	−0.9%	2.8%	0.6%
10	Japan OTC Equity Fund	−7.7%**	6.2%**	10.7%**	9.3%**
11	Malaysia Fund	3.2%**	8.1%**	8.1%**	7.4%**
12	Mexico Equity & Income Fund	−6.1%**	22.7%**	23.6%**	24.2%**
13	Mexico Fund	1.0%	13.7%**	14.3%**	15.7%**
14	New Germany Fund	19.5%**	5.5%**	5.5%**	4.3%**
15	Singapore Fund	−4.7%**	2.3%	5.5%**	9.2%**
16	Spain Fund	12.0%**	7.0%**	7.6%**	8.6%**
17	Swiss Helvetia Fund	7.4%**	7.1%**	8.2%**	9.2%**
18	United Kingdom Fund	14.6%**	4.6%**	3.4%**	2.5%**
	Average	7.6%	6.8%	7.7%	7.6%
	% Positive		83%	94%	94%
			χ^2 Statistics		
	H_0: Sum of Parameters = 0		72.275**	180.734**	348.169**
	H_0: Each Parameter = 0		213.281**	432.698**	860.429**
Panel B: Funds With No Corresponding WEBS					
1	Argentina Fund	−3.0%**	7.5%**	8.5%**	10.6%**
2	Brazil Fund	−0.7%	9.1%**	12.3%**	15.0%**
3	Brazilian Equity Fund	−4.4%**	13.0%**	18.1%**	19.8%**
4	Chile Fund	9.8%**	−1.4%	0.3%	0.5%
5	China Fund	−7.9%**	10.5%**	14.1%**	19.1%**
6	First Israel Fund	−2.7%*	16.3%**	16.1%**	16.7%**
7	Fidelity Philippine Fund	18.2%**	2.4%**	1.9%**	1.6%**
8	Greater China Fund	2.3%**	9.4%**	12.0%**	14.3%**
9	India Fund	9.5%**	−10.6%**	−6.5%**	−7.6%**
10	India Growth Fund	−6.2%**	−4.7%	−0.3%	−1.5%
11	Irish Investment Fund	12.6%**	1.8%	3.4%**	3.1%**
12	Jardine Fleming China Region Fund	0.1%	8.4%**	12.6%**	15.6%**
13	Jardine Fleming India Fund	1.0%	1.1%	2.1%	1.9%
14	Korea Equity Fund	5.0%**	−1.6%	−2.1%	−4.4%**
15	Korea Fund	−8.7%**	0.3%	−0.2%	−1.1%
16	Korean Investment Fund	3.1%**	3.1%	2.4%	0.0%
17	Morgan Stanley Indian Fund	5.4%**	−8.9%**	−5.8%**	−9.5%**
18	New South Africa Fund	18.8%**	0.9%	0.9%	0.9%
19	Pakistan Investment Fund	18.0%**	−11.3%**	−12.2%**	−15.3%**
20	ROC Taiwan Fund	−0.9%	−5.1%*	−4.4%*	4.0%**
21	Taiwan Fund	−4.5%**	−7.6%*	−3.6%	6.4%**

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Table 2 *(continued)*

Fund Number	Fund Name	Intercept	$\hat{\beta}_j$ for post-event window of:		
			13-Weeks	26-Weeks	52-Weeks
Panel B: Funds With No Corresponding WEBS					
22	Templeton China World Fund	2.9%**	11.4%**	13.0%**	13.4%**
23	Thai Capital Fund	10.2%**	−3.4%**	−4.3%**	−10.5%**
24	Thai Fund	10.7%**	−3.5%**	−5.8%**	−13.0%**
25	Turkish Investment Fund	−22.7%**	13.6%*	17.7%**	22.6%**
	Average	2.6%	2.0%	3.6%	4.1%
	% Positive		60%	60%	64%
			χ^2 Statistics		
	H_0: Sum of Parameters = 0		2.72	16.546**	41.663**
	H_0: Each Parameter = 0		453.925**	939.345**	1590.159**

*Significant at 5%; **Significant at 1%.

Results are from a systems of equations regression of weekly percentage discound against a dummy variable, D, that is constructed with ones in the post-event window. For each individual firm j, the equation is: $Y_j = \mu_j + \beta_j D + \varepsilon_j$, with $\varepsilon_j \sim N(0, \sigma^2)$.

model, which suggests either no reaction or an overreaction to an integrating event. From the economic perspective, the increase could result from either a reduction in the market segmentation premium or a migration to index funds with lower fees. However, this test does not differentiate between the two explanations.

One interesting observation from Panel A is the relative magnitudes of the β_j estimates for the three Mexican Funds. For example, in the 13-week window, the β_j estimates for the Emerging Mexico Fund, the Mexico Equity & Income Fund, and the Mexico Fund are 22.1%, 22.7%, and 13.7%, respectively. These estimates are much larger in magnitude than those of any other funds. This raises the possibility that our results are overly sensitive to changes in Mexican fund discounts. To test this possibility, we re-run the model dropping the three Mexican funds. Although not shown, the hypothesis that the individual β_j estimates are jointly zero and the hypothesis that the sum of the β_j estimates equals zero again reject at the 1% level. These results hold across all three post-event windows. Hence, we feel our results are not unduly influenced by the Mexican funds.

Table 2—Panel B examines the effect of WEBS trading on funds with no corresponding WEBS index. For the 13-week post-event window, only 9 of the 25 funds (36%) show a significantly positive (1% or 5% level) increase in mean percentage discount. Hence, over the short-term, a lower proportion of these funds experience a WEBS induced impact. Also, the average change in percentage discount is 2%, well below the 6.8% change experienced by funds with a corresponding WEBS index. Finally, the hypothesis that the sum of the β_j parameters equals zero cannot be rejected at the 5% level. For the 26-week period, 11 out of 25 funds (44%) experience a significantly positive increase in mean percentage discount evaluated at the 5% level. The average change in percentage discount is 3.6%, and the test

that the sum of the coefficients equals zero rejects at the 1% level. The 52-week run has similar results.

Table 2 suggests that WEBS trading impacts both groups of funds. However, the effect appears to be larger for funds with a corresponding WEBS index. Hence, we specifically test for a differential using the two-factor difference in means test.

4.2. Results of the two-factor test

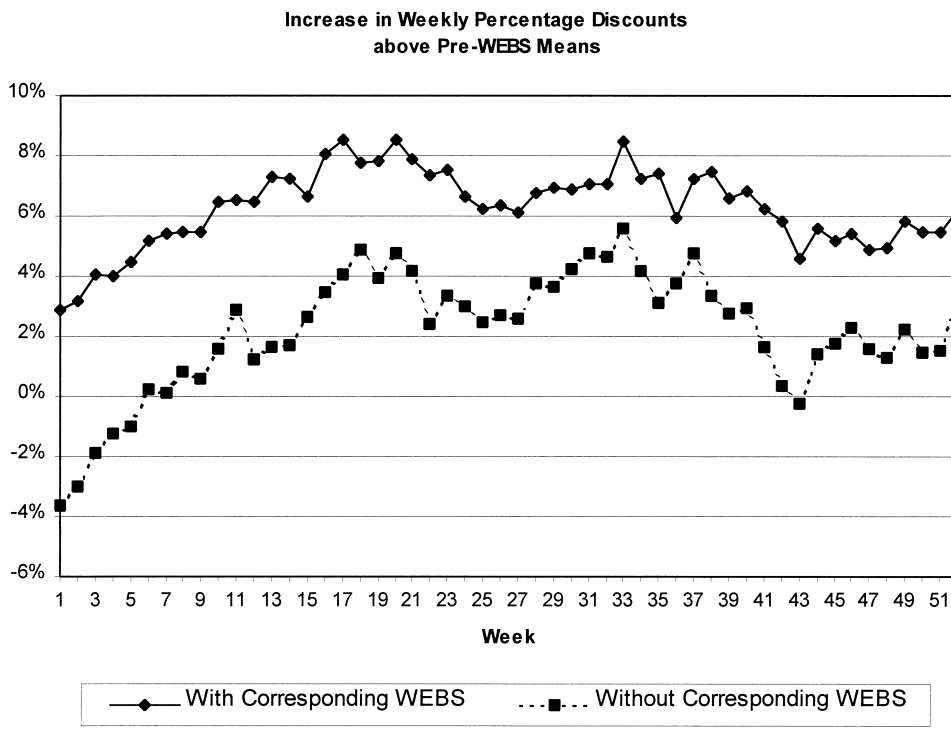
We run the two-factor model for a pooled sample of all 43 CECFs. Analogous to the systems of equations models, we run the model separately for each of the three post-event windows. Table 3—Panel A reports the parameter estimates, and Panel B presents the results of coefficient restriction tests on each of the three null hypotheses. Focusing on the 13-week post-event period, we reject the first two null hypotheses. These findings are consistent with the results from the modified event study. In particular, mean discounts for both categories of funds increase after the introduction of WEBS trading. Moreover, we reject the third null hypothesis, confirming a statistically significant differential impact. These results hold for both the intermediate and long-term windows as well. Overall, the stronger impact on corresponding closed-end funds is consistent with the economic explanation.

These results are easily observed in Chart 1 below. The chart shows the increase in average percentage discounts above pre-event means, for CECFs with and without corresponding WEBS indexes. Clearly CECFs with a corresponding WEBS index have larger and more persistent increases in percentage discounts relative to CECFs with no corresponding WEBS.

4.3. Volume analysis

Since the previous tests do not directly differentiate between a market segmentation effect and an agency effect, we also examine the impact on trading volume. If CECF agency costs are excessive, then investors should migrate to lower fee WEBS, reducing CECF volumes. As before, the impact should be more pronounced in funds with a corresponding WEBS index. To test this hypothesis, we re-run the systems of equations models with volume data replacing percentage discounts. In particular we now let Y_j be a vector of average daily volumes (observed weekly) for fund j. As before, we run the models separately by fund group and by post-event measurement window.

Table 4—Panel A presents the results of the volume analysis for funds with a corresponding WEBS index. For this model, a negative β_j coefficient indicates a decrease in trading volume after the introduction of WEBS trading. For the 13-week window, 11 of the 14 (79%) individual funds experience a decrease in mean trading volume. However, only one decrease is significant (5% level). Furthermore, we fail to reject either the null hypothesis that all funds have zero β_j parameters or that the sum of the parameters is zero. Thus in the short-term, the direction of the volume change is consistent with reduced demand, but the level is not significant. For the 26-week window, only 9 of the decreases are significant at the 1% or 5% levels. Further, the hypotheses that the parameters sum to zero and are jointly zero reject. Results of the 52-week window are similar. Compared to percentage



discounts, the change in trading volume appears to materialize more slowly and impact fewer funds.

Table 4—Panel B presents the results of the volume analysis for the 21 firms with no corresponding WEBS index. For the 13-week post-event window 13 of the 21 (62%) funds experience decreased volumes, but only 1 volume decrease is significant at even the 5% level. Also, the sum of the β_j parameters is not significantly different from zero, indicating that the introduction of WEBS trading does not have a significant effect on the trading volumes of this group. Over the 26-week window, the number of funds with lower volume remains at 13, with two of the changes significant at the 5% level. Also, the null hypothesis that the sum of the volume changes equals zero cannot be rejected. Similar results occur as we extend the window to 52 weeks. Interestingly, the null hypothesis that the β_j parameters are jointly zero easily rejects for all three event windows. This occurs because some of the volume changes are positive, some are negative, but the average change is not significant. Overall, these results suggest a negligible impact on trading volume for funds with no corresponding WEBS. As with discounts, we next examine the differential impact on each group's trading volume.

Table 5 presents the results of the two factor test on volume data. Panel B contains the results of the three hypothesis tests. For the 13-week post event window the average daily trading volume for funds with a corresponding WEBS index decreases by 14,340 shares, which is significant at the 5% level. In contrast, funds with no corresponding WEBS index show on average an insignificant increase in volume of 252 shares. However, the differential

Table 3
Two Factor Difference in Means Test (Dependent Variable = Percentage Discount)

	Post-Event Window of:		
	13 Weeks	26 Weeks	52 Weeks
Panel A: Parameter Estimates			
μ	7.2%**	7.9%**	8.0%**
α	2.2%**	2.9%**	3.0%**
γ	−3.7%**	−3.5%**	−3.4%**
θ	−1.2%**	−1.0%**	−0.9%**
Panel B: Hypotheses Test Results			
Ho			
$(\mu_{21} - \mu_{11}) = 0$	6.8%**	7.7%**	7.6%**
$(\mu_{22} - \mu_{12}) = 0$	2.1%**	3.7%**	4.2%**
$(\mu_{21} - \mu_{11}) - (\mu_{22} - \mu_{12}) = 0$	4.7%**	4.0%**	3.4%**

*Significant at 5%; **Significant at 1%.

Results are from a pooled regression of percentage discounts against three dummy variables. Letting Y be a percentage discount, the regression model is $Y = \mu + \alpha D_1 + \gamma D_2 + \Theta D_1 D_2 + \varepsilon$, where:

$$D_1 = \begin{cases} 1 \text{ if the observation is in the pre-event estimation window} \\ -1 \text{ if the observation is in the post-event window} \end{cases}$$

and

$$D_2 = \begin{cases} 1 \text{ if the observation is from a fund with a corresonding WEBS index} \\ -1 \text{ if the observation is from a fund with no corresponding WEBS Index} \end{cases}$$

The parameter μ is the average percentage discount for all funds, α measures the effect due to the time factor, γ measures the effect due to fund type, and Θ measures the interaction effect between the two factors. let the subscript i equal one for a pre-event observation and two otherwise. Let the subscript k equal one for an observation from a fund with a corresponding WEBS index and two otherwise. Then the mean percentage discount (μ_{ik}) for each level of each factor is recoverable from the regression model.

short-term volume impact is not significant. By the end of the 26-week window, the relative decrease in volume for funds with a corresponding WEBS index is significant, and this result persists into the 52-week window. These results corroborate a slowly emerging decrease in demand for CECFs with a corresponding WEBS index. This provides indirect evidence of investor migration to the lower fee WEBS indexes, which is consistent with an agency cost explanation of closed-end fund discounts. However, since the tests on percentage discounts show a more immediate and wide-spread impact, we cannot rule out a market segmentation effect.

5. Summary

This paper examines the impact of country specific index trading on closed-end country fund discounts. CECFs provide individual investors with diversified, country specific, foreign investments. However, like all closed-end funds, they often trade at a discount

Table 4
Modified Event Study Results (Dependent Variable = Average Daily Volume).

Fund Number	Fund Name	Intercept	$\hat{\beta}_j$ for post-event window of: 13-Weeks	26-Weeks	52-Weeks
Panel A: Funds with Corresponding WEBS					
1	Austria Fund	23,821**	−11,336	−8,356	−3,925
2	First Australia Fund	47,950**	−12,281	−15,152	−7,090
3	France Growth Fund	29,818**	−7,688	−9,914	−2,313
4	Germany Fund	47,193**	15,701	−19,359*	−6,464
5	Italy Fund	29,930**	547	7,493	9,625
6	Japan Equity Fund	62,249**	−37,641*	−37,418**	−23,327**
7	Japan OTC Equity Fund	52,992**	−19,615	−23,954	−16,834
8	Malaysia Fund	56,566**	−13,250	−15,285	−15,106
9	Mexico Equity & Income Fund	366,385**	−78,608	−126,024	−141,585**
10	Mexico Fund	87,098**	548	1,321	4,207
11	New Germany Fund	24,407**	−6,668	−5,168	−622
12	Singapore Fund	27,986**	−8,010	−11,679*	−1,998
13	Spain Fund	28,332**	−6,340	−10,517	−2,800
14	Swiss Helvetia Fund	34,115**	−5,815	−9,996*	−8,015*
	Average	65,632	−13,604	−20,286	−15,446
	% Negative		79%	86%	86%
				χ^2 Statistics	
	H_0: Sum of Parameters = 0		2.721	8.823	9.765**
	H_0: Each Parameter = 0		12.500	36.912**	31.007**
Panel B: Funds With No Corresponding WEBS					
1	Argentina Fund	55,172**	−1,711	−13,807	−15,536
2	Brazil Fund	89,375**	−20,4241	−20,659	−12,123
3	Brazilian Equity Fund	44,358**	15,141	13,741	4,686
4	Chile Fund	37,132**	9,802	−10,713	4,987
5	China Fund	44,346**	4,977	−11,419	−10,438
6	First Israel Fund	37,761**	15,592	2,685	6,267
7	Fidelity Philippine Fund	128,634**	3,412	8,647	7,825
8	Greater China Fund	34,954**	−4,600	−2,351	−1,083
9	India Fund	16,248**	−7,740	−7,356*	−6,256*
10	India Growth Fund	9,983**	4,871	3,390	4,521*
11	Irish Investment Fund	36,663**	−2,833	−8,106	−4,994
12	Jardine Fleming China Region Fund	114,103**	−47,334	−33,811	30,006
13	Korea Fund	29,208**	−7,203	−7,038	4,125
14	Korean Investment Fund	38,387**	−6,156	−8,002	−4,654
15	Pakistan Investment Fund	119,164**	71,651*	30,593	54,364*
16	ROC Taiwan Fund	43,872**	−12,432	−15,834*	−15,532**
17	Taiwan Fund	57,038**	−8,838	−9,188	−2,644

(continued on next page)

to NAV, which may be a concern to individual investors. Country specific indexes such as WEBS provide similar international diversification benefits without the concern over discounts.

Explanations of discounts fall into two categories: behavioral and rational economic explanations. Behavioral modes suggest either no reaction or an overreaction from investors in response to the introduction of WEBS. Rational economic models suggest

Table 4 *(continued)*

Fund Number	Fund Name	Intercept	$\hat{\beta}_j$ for post-event window of: 13-Weeks	26-Weeks	52-Weeks
Panel B: Funds with No Corresponding WEBS					
18	Templeton China World Fund	73,343**	16,103	7,295	10,695
19	Thai Capital Fund	36,819**	−14,504	8,919	8,319
20	Thai Fund	24,863**	−1,487	−2,517	3,340
21	Turkish Investment Fund	16,096**	−165	1,827	10,771
	Average	51,787	292	−3,510	3,650
	% Negative		62%	62%	43%
			χ^2 Statistics		
	H_0: Sum of Parameters = 0		0.055	1.017	1.722
	H_0: Each Parameter = 0		55.359**	66.137**	83.790**

*Significant at 5%; **Significant at 1%.

Results are from a systems of equations regression of each week's average daily volume against a dummy variable, D, that is constructed with ones in the post-event window. For each individual firm j, the equation is: $Y_j = \mu_j + \beta_j D + \varepsilon_j$, with $\varepsilon_j \sim N(0, \sigma^2)$.

Table 5

Two Factor Difference in Means Test (Dependent Variable = Average Daily Volume).

	Post-Event Window of: 13 Weeks	26 Weeks	52 Weeks
Panel A: Parameter Estimates			
μ	55,186**	52,726**	55,760**
α	−3,522	−5,947**	−2,949*
γ	−3,274	−2,734	−2,148
θ	3,648	4,187	4,774**
Panel B: Hypotheses Test Results Ho			
$(\mu_{21} - \mu_{11}) = 0$	−14,340**	−20,268**	−15,446**
$(\mu_{22} - \mu_{12}) = 0$	252	−3,520	3,650
$(\mu_{21} - \mu_{11}) - (\mu_{22} - \mu_{12}) = 0$	−14,592	−16,748*	−19,096**

*Significant at 5%; **Significant at 1%.

Results are from a pooled regression of each week's average daily volume against three dummy variables. Letting Y be a weekly observation of average daily volume, the regression model is $Y = \mu + \alpha D_1 + \gamma D_2 + \Theta D_1 D_2 + \varepsilon$, where:

$$D_1 = \begin{cases} 1 \text{ if the observation is in the pre-event estimation window} \\ -1 \text{ if the observation is in the post-event window} \end{cases}$$

and

$$D_2 = \begin{cases} 1 \text{ if the observation is from a fund with a corresonding WEBS index} \\ -1 \text{ if the observation is from a fund with no corresponding WEBS Index} \end{cases}$$

The parameter μ is the grand mean daily volume for all funds, α measures the effect due to the time factor, γ measures the effect due to fund type, and Θ measures the interaction effect between the two factors. Let the subscript i equal one for a pre-event observation and two otherwise. Let the subscript k equal one for an observation from a fund with a corresponding WEBS index and two otherwise. Then the grand mean daily volume (μ_{ik}) for each level of each factor is recoverable from the regression model.

that discounts will increase for CECFs with a corresponding WEBS index due to increased competition and reduced market segmentation premiums.

We use a modified event study to test the impact of country specific index trading on CECF discounts. Our results are consistent with a rational economic explanation of the discount. In particular, the percentage discount increases after the introduction of WEBS trading. The increase is stronger for CECFs with a corresponding WEBS index and remains permanent over a 52-week interval. Funds with a corresponding WEBS index also experience a decrease in trading volume, further suggesting a migration away from CECFs.

Our results are particularly important to individual investors. Country specific indexes, such as WEBS help investors optimize risk-return tradeoffs by lowering the costs of international diversification. Competitive pressures force CECF managers to improve performance and lower agency costs, or face additional migratory pressures. Nevertheless, a caution to individual investors is in order. Our results suggest that CECF investors should be wary of widening discounts for funds facing the introduction of a competing country specific index.

Appendix A
Closed-end Country Funds by Group

CECFs with a Corresponding WEBS Index		CECFs without a Corresponding WEBS Index
Fund Name and Symbol	WEBS Index and Symbol	Fund Name and Symbol
Austria Fund (OST)	Austria (EWO)	Argentina Fund (AF)
First Australia Fund (IAF)	Australia (EWA)	Brazil Fund (BZF)
Canadian General Investors (CGI)	Canada (EWC)	Brazilian Equity Fund (BZL)
France Growth Fund (FRF)	France (EWQ)	Chile Fund (CH)
Germany Fund (GER)	Germany (EWG)	China Fund (CHN)
New Germany Fund (GF)	Germany (EWG)	First Israel Fund (ISL)
Italy Fund (ITA)	Italy (EWI)	First Philippine Fund (FPF)
Japan Equity Fund (JEQ)	Japan (EWJ)	Greater China Fund (GCH)
Japan OTC Fund (JOF)	Japan (EWJ)	India Fund (IFN)
Mexico Equity & Income Fund (MXE)	Mexico (EWW)	India Growth Fund (IGF)
Emerging Mexico Fund (MEF)	Mexico (EWW)	Irish Investment Fund (IFL)
Mexico Fund (MXF)	Mexico (EWW)	Jardine Fleming China Region Fund (JFC)
Singapore Fund (SGF)	Singapore (EWS)	Jardine Fleming India Fund (JFI)
Spain Fund (SNF)	Spain (EWP)	Korea Equity Fund (KEF)
Growth Fund of Spain (GSP)	Spain (EWP)	Korea Fund (KF)
Swiss Helvetia Fund (SWZ)	Switzerland (EWL)	Korean Investment Fund (KIF)
United Kingdom Fund (UKM)	United Kingdom (EWU)	Morgan Stanley India Fund (IIF)
		New South Africa Fund (NSA)
		Pakistan Investment Fund (PKF)
		ROC Taiwan Fund (ROC)
		Taiwan Fund (TWN)
		Templeton China World Fund (TCH)
		Thai Capital Fund (TC)
		Thai Fund (TTF)
		Turkish Investment Fund (TKF)

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