Global versus Local Asset Pricing: Evidence from Arbitrage of the MSCI Index Change

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Abstract

Large-scale simultaneous asset demand shocks like index revisions modify stock betas market-wide and generate testable cross sectional asset pricing implications. This paper develops a model of limited arbitrage which characterizes the cross-sectional return dynamics around a partially anticipated public announcement of an index revision. Arbitrage by risk averse speculators implies that stock returns prior to the announcement are not only positively proportional to the anticipated beta decrease of each stock, but also negatively proportional to the marginal arbitrage risk contribution of each speculative position. The redefinition of the MSCI international equity index in 2001 and 2002 provides a powerful event study to test these predictions and delivers strong evidence in favor of the new model. Importantly, the global nature of the MSCI index revision implies that global and local beta changes differ substantially along with the respective marginal arbitrage risk incurred to exploit them. Testing which beta changes and marginal arbitrage risk terms are price relevant reveals that MSCI stocks are priced globally and not locally.

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1 Introduction

This paper explores the security price dynamics for an event in which a large number of stocks worldwide experience changes in the investor demand. In financial markets, demand shocks often affect more than one security and the size and direction of the demand change may differ across securities. Fund managers may for example liquidate proportionally to their holdings when faced with large fund outflows. Simultaneous demand shocks may also result from the build-up or liquidation of hedging positions. Finally, they occur (as in this study) when equity indices are redefined with new stocks included and other stocks deleted from the index.

Exogenous multi-asset demand shocks like large-scale index revisions are of particular interest, because they modify stock betas market-wide and therefore generate testable cross sectional asset pricing implications. The intuition is straightforward. Stocks for which the weight change is positive experience an effective asset supply contraction after accounting for the increased demand from index tracking investors. A lower residual supply for up-weighted stocks will decrease the risk contribution to total market risk of all stocks which have a strong positive covariance with the up-weighted stocks. In analogy to the intuition of the Capital Asset Pricing Model (CAPM), a lower risk contribution to the total tradeable asset supply risk implies a lower stock "beta" and earns a lower stock specific expected return. A multi-asset demand shock like an index revision represents a simultaneous shock to all stock betas. More interesting still, such a shock generally modifies the global stock beta differently than the local stock beta. Hence, one can interpret a global equity index change like MSCI index redefinition as a natural experiment on global versus local asset pricing. An important contribution of this paper is to show that MSCI stocks are priced globally and not locally: global and not local premium changes explain the cross section of price changes around the public announcement of the index revision.

The current paper makes two theoretical contributions. First, it generalizes the existing theory of limited arbitrage under multi-asset demand shocks.[1] The new framework allows for asymmetric information where arbitrageurs learn about an imminent index revision prior to its public announcement. Arbitrageurs can acquire net positions against uninformed liquidity

[1] The multi-asset framework developed in Greenwood (2005) has obvious shortcomings. Most importantly, all investors are assumed to have identical information and represent arbitrageurs. But homogeneity among (non-index tracking) investors implies that the equilibrium price adjustment to the demand shock occurs without any speculative position taking as the net asset supply to the arbitrageurs is fixed. Arbitrage profits are therefore zero. Moreover, the absence of speculative positions means that any hedging or risk management motive for these positions disappears and their corresponding price impact is also discarded. Such a minimalist arbitrage framework is very detached from a practitioner's notion of arbitrage to which asymmetric information is essential.

suppliers prior to the announcement in order to make arbitrage profits. It is shown that the optimal arbitrage position features two distinct terms. It has a 'return seeking component' proportional to the beta change of each stock. Moreover, it features a 'hedge component' proportional to the stock-specific marginal arbitrage risk contribution. Both dimensions of the optimal arbitrage position codetermine the cross-sectional return pattern prior to the public announcement, i.e. the pre-announcement returns. The stock-specific arbitrage risk can be approximated by the product of the *squared* return covariance matrix ($\Sigma\Sigma$) and the vector of weight changes from old to new index weights $w^n - w^o$. Intuitively, the return seeking component of the arbitrage position is proportional to the beta change $\Sigma(w^n - w^o)$ and generates an absolute arbitrage risk $(w^n - w^o)'\Sigma\Sigma\Sigma(w^n - w^o)$ and a marginal arbitrage risk $\Sigma\Sigma(w^n - w^o)$. Optimization in the mean-variance space requires arbitrageurs to choose a portfolio which optimizes the trade-off between expected arbitrage returns and marginal arbitrage risk in each stock. The optimal arbitrage strategy down-weights stocks with a positive marginal arbitrage risk contribution and up-weights stocks with a negative marginal arbitrage risk contribution. The marginal arbitrage risk thus becomes a pricing factor along with the premium or beta change itself. The generalized model nests the Greenwood framework as a special case where stock-specific liquidity suppliers disappear. In the latter case, price adjustment occurs without any speculative position taking or arbitrage risk.

Second, international asset demand shocks such as global index revisions are shown to provide straightforward tests as to whether stocks are priced globally or locally. The global covariance matrix Σ^G of all index stocks can be decomposed into a matrix Σ^L featuring non-zero covariance elements only between stocks in the same country and a complementary matrix $\Sigma^{Int} = \Sigma^G - \Sigma^L$ featuring only cross country stock covariances (and zeros otherwise). The global beta changes are proportional to $\Sigma^G(w^n - w^o)$, while local beta change are proportional to $\Sigma^L(w^n - w^o)$. Global asset markets are segmented if the difference between the global and local beta change given by $\Sigma^{Int}(w^n - w^o)$ does not help to explain event returns. Alternatively, market integration with respect to asset pricing implies that local beta changes $\Sigma^L(w^n - w^o)$ and the complementary international beta changes $\Sigma^{Int}(w^n - w^o)$ feature the same quantitative influence on event returns. A similar decomposition into a local and a complementary international component can also be applied to the arbitrage risk $\Sigma\Sigma(w^n - w^o)$, which allows additional inference about the degree of market integration. Interestingly, these simple asset pricing tests do not rely on the correct identification of the global or local market benchmarks. The true global or local market weights may not be known (Rolls (1977)). Asset pricing tests based on exogenous demand shocks are predicated on the correct identification of index weight changes, but not on any identification of the market itself.[2]

The empirical analysis focuses on the revision of the global MSCI index announced in December 2000 and implemented in two steps in November 2001 and May 2002. This choice has a number of advantages over the events used in previous studies. First, the weight revision concerned a total of 2566 stocks in 50 countries. It therefore presents an index change of unprecedented scope, which provides great cross-sectional power to discriminate between different theories of imperfect arbitrage and their asset pricing implications. Second, the announcement of the MSCI index revision and its implementation are separated by at least 12 months. The pre- and post-announcement event windows can therefore be easily separated from the implementation event. In the Nikkei 225 revision considered by Greenwood (2005), announcement and implementation are separated only by one week and the empirical analysis does not attempt to isolate any pre- or post-announcement effects. Third, the international dimension of the index change allows us to infer the degree of market integration with respect to asset pricing. Previous empirical work on the degree of international equity market integration has used capital market liberalization as the identifying event to measure risk premium changes (Chari and Henry (2004)). In a similar spirit, I test whether the local or international components of risk premium changes and arbitrage risk determine returns over a more sharply defined event window. Moreover, the index change in this paper is certainly more exogenous than a liberalization policy which may also correlate with changing company cash flows.

The empirical results concern both the generalized theory of arbitrage and the issue of international market integration. These can be summarized as follows:

- Pre-announcement returns are determined (positively) by the risk premium (or beta) decrease of each stock, and (negatively) by stock-specific arbitrage risk contribution of the speculative positions held against the liquidity providers. In the run-up to the (partially anticipated) announcement, speculators acquire stocks with high expected price increases (those with risk premium or beta decreases), but simultaneously hedge their speculative risk by shorting stocks which have a high marginal arbitrage risk.

- Post-announcement returns show a positive cross-sectional relationship to marginal arbitrage risk. The post-announcement price effects comes from the liquidation of the specu-

[2] None of the conclusions therefore depends on whether the MSCI index represents the global market benchmark or not. We only assume that the weight changes represent a *change* to that global market benchmark, whatever it may be.

lators' hedging positions after the index revision is known by all market participants and prices have adjusted to the beta change. The post-announcement returns show that the nested Greenwood model is strongly rejected by the data in favor of the general framework.

- The decomposition of the global premium change into local and complementary international premium changes shows that both are equally price relevant for the pre-announcement returns. The index weight increase of a Japanese stock for example alters the IBM stock price in the same magnitude as the identical index weight increase of a U.S. company if both stocks feature the same covariance with the IBM stock return.

- The decomposition of the arbitrage risk factor into a local and complementary international component confirms this conclusion for both the pre- and post-accouncement returns. Overall, the hypothesis of global asset pricing is maintained and the hypothesis of local asset pricing rejected by the data.

The finance literature includes a number of studies on the stock price impact of index inclusions and exclusions. Initially, these event studies all focus on individual price movements with overwhelming evidence that index inclusions increase share prices and exclusions decrease them.[3] The portfolio approach in Greenwood (2005) differs from these previous contributions in its focus on cross sectional asset pricing implications. A broader literature on 'liquidity effects' assesses whether demand shocks correlate with individual stock price returns. Time series studies on block purchases and sales of stocks, as well as the trades of institutional investors, have consistently uncovered evidence of temporary price pressure on individual securities conditional upon unusual demand (Lakonishok, Shleifer and Vishny (1991,1992), Chan and Lakonishok (1993, 1995)). In the international finance literature, Froot, O'Connell and Seasholes (1998) have shown that local stock prices are sensitive to international investor flows, and that transitory inflows have a positive future impact on returns. Focusing on mutual funds, Warther (1995) and Zheng (1999) have documented that investor demand effects may aggregate to the level of the stock market itself. Goetzmann and Massa (2002) show that, at daily frequency, inflows into S&P500 index funds have a direct impact on the stocks that are part of the index. This literature is generally concerned with the mere existence of liquidity effects without strong asset pricing foundations.

[3] See for example (Garry and Goetzmann (1986), Harris and Gurel (1986), Shleifer (1986), Dillon and Johnson (1991), Beniesh and Whaley (1996), Lynch and Mendenhall (1997)). Kaul et al. (2000) examine index reweighting for stocks in the Toronto Stock Exchange 300 index and find that upweighted stocks experience a persistent positive price effect. See also Denis et al. (2003) and Hedge and McDermott (2003).

Karolyi and Stulz (2003) provide a literature survey on the issue of global versus local asset pricing. Such research has increasingly tested the world CAPM in a conditional setting with time-varying expected returns, variances and covariances as exemplified in Harvey (1991); Chan, Karolyi and Stulz (1992); and De Santis and Gerard (1997). But inference here is sensitive to the GARCH specification and less restrictive specifications (with a large number of parameters) suffer from low power and only weak empirical support for the CAPM model. An additional concern is that financial globalization has modified the degree of international market integration over the two recent decades. The event study approach in this paper provides a unique snap shot on market integration using relatively recent data in a very short time window. Moreover, the cross-sectional test of market integration does not rely on the correct identification of the market benchmarks on which intertemporal models are predicated.

The paper proceeds as follows: Section 2 outlines a new model of multi-asset arbitrage. The cross-sectional stock return implications for the pre-announcement, and post-announcement periods are highlighted in propositions 1 and 2, respectively. The stock price reaction for the implementation dates is summarized in proposition 3. Proposition 4 characterizes the pre- and post-announcement return under the polar cases of complete international market integration and segmentation. Section 3 describes the MSCI index redefinition and discusses summary statistics about the index weight changes, the risk premium changes and the arbitrage risk for individual stocks. Section 3 also characterizes the total portfolio risk of the optimal arbitrage strategy relative to a passive holding strategy in the old MSCI index. Section 4 provides the evidence on the pre- and post-announcement effect, the implementation effect and the degree of global versus local asset price determination. Section 5 concludes.

2 Theory and Hypotheses

2.1 Model Assumptions

This section develops a simple limits-to-arbitrage model which allows me to analyze the return effects of demand shocks in a multi-asset market setting. A set of N financial assets is priced over 4 periods. The market characteristics are summarized in Assumption 1:

Assumption 1: Market Structure, Asset Supply and Liquidation Value

The financial market with N risky assets allows trading at 4 different times $t = 0, 1, 2, 3$. In period $t = 4$ all assets are liquidated at liquidation prices given by

$$p_3 = \mathbf{1} + \sum_{t=1}^{4} \varepsilon_t.$$

where ε_t denotes serially uncorrelated mean zero innovation learned by all market participants at time t. The covariance of innovations ε is given by the matrix Σ. The asset supply in periods $t = 0, 1, 2$ is given by S. In period $t = 3$, a demand shock reduces the asset supply to $S - u$, where $u = w^n - w^o$ represents the exogenous demand change from old index weights w^o to new index weights w^n. The ex ante $(t = 0)$ expected liquidation price is normalized to the unit vector $\mathbf{1}$.

The stochastic liquidation value generates asset investment risk. The index revision is modeled like in Greenwood (2005) as an exogenous change in the asset supply. Stocks with increased weight face a higher demand by index tracking funds so that their net asset supply $S - u$ is reduced. The demand shock u from the index investors is completely price inelastic. Index investors therefore do not qualify as counterparty to intertemporal arbitrage trades. The behavior of the index investors is fully captured by the one-time demand shock.

A new feature of the proposed model is the introduction of liquidity supplying agents. These are the potential counterparties to the arbitrageurs seeking a net arbitrage position. The arbitrage opportunity is further imbedded in the assumption that liquidity suppliers learn about the exogenous liquidity shock only with a delay of one period. It is then shown that the existence of less informed liquidity suppliers significantly modifies the cross-sectional price patterns of event returns. Assumption 2 characterizes the investment behavior of these two types of market participants:

Assumption 2: Arbitrageurs and Linear Liquidity Supply

A unit interval of market participants can be grouped into a set $[0, \lambda]$ of risk arbitrageurs and a set of liquidity suppliers $[\lambda, 1]$. Arbitrageurs have a CARA utility and a risk aversion parameter ρ, and access to a riskless asset of zero return. Their optimal demand vector of a CARA investor follows as

$$x^A = (\rho\Sigma)^{-1}\widetilde{\mathcal{E}}_t(p_{t+1} - p_t),$$

where p_t denotes the price vector in period t and $\widetilde{\mathcal{E}}_t$ their expectation for the consecutive price appreciation. Liquidity suppliers provide in each stock a linear asset supply which depends on the asset supply elasticity γ and is given by the vector

$$x^S = \gamma\overline{\mathcal{E}}_t(p_{t+1} - p_t),$$

where $\overline{\mathcal{E}}_t$ characterizes the expectations of the liquidity suppliers.

The arbitrageurs are optimizing agents who maximize the CARA utility over their one period investment horizon. The liquidity suppliers by contrast represent an ad hoc addition to the model. Representative agent model appear generally inconsistent with existing evidence for steep demand curves for individual stocks (Petajisto (2008)). Limited market participation and short-term liquidity supply by financial intermediaries (market makers) on a stock by stock base are plausible assumptions to explain low supply elasticity of stocks. The linear liquidity supply formulated in assumption 2 is best interpreted as a short-cut to capture such market friction. The Greenwood framework is nested in the specification and recovered for a parameter $\lambda = 1$ when only arbitrageurs constitute the market.[4]

An apparently restrictive assumption consists of imposing an identical parameter γ for the liquidity supply elasticity upon all stocks. It is straightforward to relax this assumption. The scalar γ can be replaced by a matrix

$$\Gamma = \begin{bmatrix} \gamma_1 & 0 & \cdots & 0 \\ 0 & \gamma_2 & & \\ \vdots & & \ddots & \\ 0 & & & \gamma_n \end{bmatrix},$$

where stock specific liquidity supply elasticities feature as the diagonal elements. None of the model insights depend on this modification.[5] The empirical section generally abstracts from liquidity differences across stocks and assumes that such differences average out in the cross-sectional regressions.

The very existence of arbitrage opportunities also depends on information asymmetries between different market participants. This feature is incorporated by assuming that the arbitrageurs learn about the index weight change in period $t = 1$, but that liquidity suppliers learn about it only in period $t = 2$. This allows arbitrageurs to exploit their information advantage in period $t = 1$ with interesting testable cross-sectional asset pricing effects. A second new assumption is that arbitrageurs and (one period later) liquidity suppliers estimate the magnitude

[4] Formally, Greenwood builds on the asset pricing framework in Hong and Stein (1999) and Barberis and Shleifer (2003) and assumes a time varying dividend process. I dispense with the dividend process and just assume a stochastic liquidation value. No important insight is lost under this simplification.

[5] It can be shown that stock-sepcific liquidity differences do not alter the return effect of the premium change, which is still proportional to Σu. Intuitively, arbitrageurs modifiy their speculative demand so as to equalize the price impact of their demand across stocks with different liquidity. However, the arbitrage risk factor differs across stocks of different liquidity since lower speculative positions for low liquidity stocks require also smaller hedge positions. The arbitrage risk effect on returns is proportional to $\Gamma\Sigma\Sigma u$.

of the demand shock only with an error. In particular, it is assumed that the expected demand shock $\widetilde{\mathcal{E}}_1(u) = \widetilde{\mathcal{E}}_2(u) = \overline{\mathcal{E}}_2(u) = \widehat{u} = ku$ of both arbitrageurs and liquidity suppliers deviates from true shock u by a scalar factor $k > 0$. By assumption, the estimation error $u - \widehat{u} = (1-k)u$ affects all stocks in the same direction and in equal proportion, which means that it concerns only the overall magnitude of the shock. The true magnitude of the demand shock is learnt only upon implementation of the index change at time $t = 3$. Thus, for the case $k \neq 1$, additional price effects are obtained when the implementation of the index change occurs. Assumption 3 summarizes the information structure:

Assumption 3: Information Structure for the 4 Trading Events

At time $t = 0$, arbitrageurs and liquidity suppliers know that the expected liquidation value of all assets is one and assume a constant supply vector S.

At time $t = 1$, only the arbitrageurs lean about the exogenous demand shock $u = w^n - w^o$ prior to the announcement of the index change. The arbitrageurs over- or underestimate its magnitude by a factor k so that they hold beliefs $\widehat{u} = \widetilde{\mathcal{E}}_1(u) = ku$ where $k \approx 1$.

At time $t = 2$, the liquidity suppliers also learn about demand shock and then share the beliefs of the arbitrageurs, hence $\overline{\mathcal{E}}_2(u) = \widehat{u}$.

At time $t = 3$, the exogenous demand shock occurs. At that moment arbitrageurs and liquidity suppliers learn the true magnitude of the demand shock so that $\widetilde{\mathcal{E}}_3(u) = \overline{\mathcal{E}}_3(u) = u$.

For simplicity, stock price changes at time $t = 1, 2, 3$ are referred to as *pre-announcement effect*, *post-announcement effect* and *implementation effect*, respectively. For the special case that $k = 1$, both arbitrageurs and liquidity providers correctly anticipate the magnitude of the demand shock. In the latter case, no specific cross-sectional return pattern is predicted for the implementation event as shown in the next section. But in practice, it may be difficult to predict the exact magnitude of a demand shock. For example, in the case of the MSCI index redefinition, the exact global capitalization of all MSCI index funds was unknown.[6]

2.2 Model Solution and Hypothesis

It is straightforward to solve the model backwards period by period. The CARA utility assumption for the arbitrageurs and the linear liquidity supply result in a linear asset demand for all

[6] Even MSCI itself seems to dispose of very vague estimates of this capitalization.

stocks. Market clearing then implies

$$\begin{aligned} S &= \lambda(\rho\Sigma)^{-1}\widetilde{\mathcal{E}}_t(p_{t+1}-p_t) + (1-\lambda)\gamma\overline{\mathcal{E}}_t(p_{t+1}-p_t) \quad \text{for} \quad t=0,1,2 \\ S-u &= \lambda(\rho\Sigma)^{-1}\widetilde{\mathcal{E}}_t(p_{t+1}-p_t) + (1-\lambda)\gamma\overline{\mathcal{E}}_t(p_{t+1}-p_t) \quad \text{for} \quad t=3. \end{aligned}$$

The index change by $u = w^n - w^o$ is represented as an asset demand shock at time $t = 3$. Up-weighted stocks are held to a larger extent by index funds and this reduces the residual supply in these stocks. Unlike the liquidity suppliers, arbitrageurs anticipate this price change at time $t = 1$ and exploit their information advantage. They accumulate net speculative positions under the belief that $\widetilde{\mathcal{E}}_1(u) = \widehat{u}$, whereas the lack of such information implies $\overline{\mathcal{E}}_1(u) = 0$ for the liquidity suppliers. After the public announcement of the index change at time $t = 2$, both the arbitrageurs and the liquidity suppliers anticipate the demand shock with $\widetilde{\mathcal{E}}_2(u) = \overline{\mathcal{E}}_2(u) = \widehat{u}$.[7] Finally, the exact magnitude of the shock becomes known at the implementation date of the index change given by time $t = 3$, when $\widetilde{\mathcal{E}}_3(u) = \overline{\mathcal{E}}_3(u) = u$.

In order to gain intuition for the different price effects along the time line, I first consider the baseline model with no liquidity supply ($\lambda = 1$) and correct expectations about the magnitude of the demand shock ($k = 1$). The (adjusted) price $\overline{p}_t^j$ for stock j is defined as the equilibrium price p_t^j adjusted for the sum of successive innovations to the liquidation price, that is

$$\overline{p}_t^j = p_t^j - \sum_{i=1}^{t} \varepsilon_i^j.$$

The terminal value is (by construction) given by $\overline{p}_4^j = 1$. In the absence of any demand shock, the period liquidity premium is simply $r^j = \frac{\rho}{\lambda}[\Sigma S]_j$, which implies an initial stock price 4 periods earlier of $\overline{p}_0^j = 1 - 4r^j$. The demand shock u changes the premium for the last period to $r_4^j = \frac{\rho}{\lambda}[\Sigma(S-u)]_j$, while all other expected returns remain unchanged. When arbitrageurs learn about the demand shock at time $t = 1$, the stock price changes by

$$\Delta\overline{p}_1^j = r^j - r_4^j = \frac{\rho}{\lambda}[\Sigma u]_j.$$

The price change $\Delta\overline{p}_1^j$ for stock j depends on the entire vector of covariances $\Sigma_{j\bullet}$ with all other stocks as well as the entire vector of weight changes u. The vector product can be positive or negative independently of the sign of the element u_j. Both up-weighted or down-weighted stocks

[7] The liquidity suppliers do not learn from the demand of the arbitrageurs in period 1 as is the case in a rational expectation equilibrium. Incorporating such learning obviously attenuates the possibility of arbitrageurs to acquire profitable speculative positions depending on the 'noise level' of other trading activity. In the limit case where the price is fully revealling, no speculative position can be acquired, which is similar to the non-existance of liquidity providers or $\lambda = 1$.

may therefore experience either a positive or negative price effect. Hence, the model does not imply any tight link between the weight change u_j of stock j and its price change at time $t = 1$. The demand shock itself at $t = 3$ does not create a price jump for the adjusted price $\overline{p}_3^j$. This baseline case represents a simplified version of the Greenwood model (2005).

The first model extension incorporates an expectational error ($k \neq 1$) with respect to the magnitude of the demand shock, but still excludes liquidity providers from the market ($\lambda = 1$). The expected return premium in the last period is now given by $\widehat{r}_4^j = \frac{\rho}{\lambda}[\Sigma(S - \widehat{u})]_j$ and the price change at time $t = 1$ follows as

$$\Delta\overline{p}_1^j = r^j - \widehat{r}_4^j = \frac{\rho}{\lambda}[\Sigma\widehat{u}]_j = \frac{\rho}{\lambda}k[\Sigma u]_j.$$

At time $t = 3$, arbitrageurs learn the true u and this triggers an additional price change

$$\Delta\overline{p}_3^j = -(r_4^j - \widehat{r}_4^j) = \frac{\rho}{\lambda}[\Sigma(u - \widehat{u})]_j = \frac{\rho}{\lambda}(1 - k)[\Sigma u]_j.$$

The larger the expectational error $(1 - k)$ about the magnitude of the demand shock, the larger the price adjustment is around the implementation event $t = 3$. Note also that no systematic price change occurs at time $t = 2$, since by assumption only symmetrically informed arbitrageurs exist and the full price adjustment (apart from the expectational error) has occurred at time $t = 1$.

The second model extension consists of a price elastic liquidity supply with $\lambda < 1$. Now, the model features less informed liquidity suppliers against which the better informed arbitrageurs can acquire speculative positions. Given a liquidation value of $\overline{p}_4^j = 1$, the price at time $t = 3$ can be determined as $\overline{p}_3^j = 1 - r_4$, where the period 4 risk premium follows as $r_4 = \left[I + (1 - \lambda)\gamma\frac{\rho}{\lambda}\Sigma\right]^{-1}\frac{\rho}{\lambda}\Sigma(S - u)]$. Without the demand shock, the equity premium is given by $r = \left[I + (1 - \lambda)\gamma\frac{\rho}{\lambda}\Sigma\right]^{-1}\frac{\rho}{\lambda}\Sigma S$. As shown in the Appendix, the pre-announcement price change now generalizes to

$$\begin{aligned}\Delta\overline{p}_1^j &= \left[\left[I + (1 - \lambda)\gamma\frac{\rho}{\lambda}\Sigma\right]^{-1}(r - \widehat{r}_4)\right]_j = \left[\left[I + (1 - \lambda)\gamma\frac{\rho}{\lambda}\Sigma\right]^{-2}\frac{\rho}{\lambda}\Sigma\widehat{u}\right]_j \\ &\approx \frac{\rho}{\lambda}k[\Sigma u]_j - 2(1 - \lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2 k[\Sigma\Sigma u]_j,\end{aligned}$$

where a linear approximation to the inverse function is used. The price or return effect at $t = 1$ is now not only determined by the change in the equity risk premium Σu, but also by the arbitrage risk term $\Sigma\Sigma u$ of the arbitrageurs.[8] Proposition 1 summarizes the return effect in the general

[8] The normalization of the liquidation price vector to **1** implies that any price change translates into an equally large event return. The market capitalizations of all stock influence the level of all risk premia, but index weight changes are the sufficient metric to capture cross-sectionally the risk premium changes.

case with liquidity providers:

Proposition 1: Pre-Announcement Returns

Prior to announcement of the weight change from old weights w^o to new weights w^n, the event return is positively proportional to the (expected) premium change $\Sigma(w^n - w^o)$ and negatively proportional to the (expected) arbitrage risk $\Sigma\Sigma(w^n - w^o)$, where Σ represents the covariance matrix of asset returns. Formally, the following linear approximation is obtained:

$$\Delta p_1 \approx \alpha_1 \times \Sigma(w^n - w^o) + \beta_1 \times \Sigma\Sigma(w^n - w^o),$$

with $\alpha_1 = \frac{\rho}{\lambda} k > 0$ *and* $\beta_1 = -2(1-\lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2 k < 0$.

Proof: See Appendix.

In the baseline case of the Greenwood model with $\lambda = 1$, the announcement price effect simplifies to the single term $\Sigma(w^n - w^o)$. This price effect represents the change in the stock specific risk contribution to the total market risk under the asset demand change $w^n - w^o$. This term is referred to as the risk premium change since it is proportional to the beta change of a stock given by $\Delta beta = -(\sigma_m^2)^{-1}\Sigma(w^n - w^o)$, where σ_m^2 represents the market variance. The index revision is therefore a large-scale modification of all stock betas and should change all stock prices proportionally given that stock cash flows remain unchanged.[9] It is also interesting to note that the price effect induced by the beta change does not depend on a correct specification of the overall asset supply S or the market benchmark, but only on the change $w^n - w^o$ of this supply. The pricing inference expressed in proposition 1 is therefore immune to the so-called Roll's critique according to which S is difficult to identify.

In the general case when $\lambda < 1$, arbitrageurs take positions to exploit their knowledge about the expected premium change $\Sigma(w^n - w^o)$ in their trading against the uninformed liquidity suppliers. The information advantage allows them to exceed the CAPM-based fair risk compensation. Optimization in the mean-variance space consists for arbitrageurs in a portfolio choice which linearly combines a 'return seeking' position with a risk reducing 'hedge' position. The 'return seeking' is best achieved by a portfolio proportional to the premium or beta change, namely $\Sigma(w^n - w^o)$. To understand the hedging position, it is useful to calculate the absolute portfolio risk of the return seeking position as $(w^n - w^o)'\Sigma\Sigma\Sigma(w^n - w^o)$. The marginal arbitrage risk of such a position follows as $\Sigma\Sigma(w^n - w^o)$. The optimal hedge position is designed to partially reverse these marginal risk contributions. A hedge portfolio $-\Sigma\Sigma(w^n - w^o)$ reduces weights in stocks

[9] The inference abstracts from the fact that growth stocks and highly leveraged stocks might be somewhat more sensitive to changes in their beta.

with positive marginal arbitrage risk contributions and increases weights in stocks with negative marginal risk contributions. An optimal arbitrage portfolio combines the 'return seeking' component and the 'risk reducing' hedge component and therefore features two distinct cross-sectional price effects characterized by the linear combination $\alpha_1\Sigma(w^n - w^o) + \beta_1\Sigma\Sigma(w^n - w^o)$ with coefficients $\alpha_1 > 0$ and $\beta_1 < 0$, respectively. A demand shock like the MSCI revision allows us to test these parameter restrictions. The corresponding empirical results are presented in section 4.1.

Next, I discuss the price behavior at time $t = 2$. After the announcement of the index change, liquidity suppliers also learn about the demand shock $\widehat{u}$. Shared beliefs between arbitrageurs and liquidity suppliers then imply a rebalancing of the stock holdings towards the market portfolio. The arbitrageurs sell their long positions in stocks with a risk premium decrease and buy back stocks with high arbitrage risk in which they hold short positions (relative to the market portfolio). Since the price adjustment in the premium change has already occurred due to the price pressure of speculative buying, the only remaining price adjustment comes from the liquidation of the risk hedging positions. Proposition 2 formalizes this intuition as follows:

Proposition 2: Post-Announcement Returns

When the liquidity suppliers learn about the weight change from old weights w^o to new weights w^n, the event return is positively proportional to the (expected) arbitrage risk $\Sigma\Sigma(w^n - w^o)$, where Σ represents the covariance matrix of asset returns. Formally, the following linear approximation is obtained:

$$\Delta p_2 \approx \beta_2 \times \Sigma\Sigma(w^n - w^o),$$

with $\beta_2 = -\beta_1 = 2(1-\lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2 k > 0.$

Proof: See Appendix.

The combination of Propositions 1 and 2 implies a distinct price pattern for the portfolio composed of weights given by the marginal arbitrage risk $\Sigma\Sigma(w^n - w^o)$. Such a portfolio can be constructed as a self-financing portfolio with zero sum of weights and a sum of weights normalized to one. This portfolio is referred to as the hedge portfolio ϖ^H since positions $-\theta\varpi^H$ constitute the optimal cross-sectional hedge against the arbitrage risk. The hedge portfolio should feature a V-shape around the announcement date of the index change. Initially, the arbitrageurs short this portfolio. This leads to an initial price decrease in the value of the portfolio captured by $\beta_1 < 0$ in proposition 1. After the announcement, the short positions are liquidated and the buying back of the stocks with weights implies a reverse up-swing in the portfolio value captured by $\beta_2 = -\beta_1 > 0$ in proposition 2. The corresponding empirical results about the post-announcement return behavior are discussed in section 4.2.

Finally, return effects at time $t = 3$ need to be discussed. It is plausible to assume that the exact magnitude of the demand shock is revealed around the implementation date when the market experiences the true demand shock. The previous pricing error due to incorrect expectations ($k \neq 1$) is reversed with an additional price adjustment proportional to $1 - k$. Proposition 3 characterizes the implementation returns as follows:

Proposition 3: Implementation Returns

On implementation of the weight change from old weights w^o to new weights w^n the return vector is proportional to unexpected premium change $(1-k)\Sigma(w^n - w^o)$ and the unexpected arbitrage risk $(1-k)\Sigma\Sigma(w^n - w^o)$, where $(1-k)(w^n - w^o) = u - \mathcal{E}(u)$ represents the prediction error for the demand shock. Formally, the following linear approximation is obtained:

$$\Delta p_3 \approx \alpha_3 \times \Sigma(w^n - w^o) + \beta_3 \times \Sigma\Sigma(w^n - w^o),$$

with $\alpha_3 = \frac{\varrho}{\lambda}(1-k)$ *and* $\beta_3 = -(1-\lambda)\gamma\left(\frac{\varrho}{\lambda}\right)^2(1-k)$. Therefore, I predict

$$\begin{array}{llllll} (i) & \alpha_3 < 0 & \textit{and} & \beta_3 > 0 & \textit{and} & k > 1 \\ (ii) & \alpha_3 = 0 & \textit{and} & \beta_3 = 0 & \textit{and} & k = 1 \\ (iii) & \alpha_3 > 0 & \textit{and} & \beta_3 < 0 & \textit{and} & 0 < k < 1, \end{array}$$

for (i) overestimation or (ii) correct estimation or (iii) underestimation of the demand shock $u = w^n - w^o$, respectively.

Proof: See Appendix.

The testable model restriction here is whether the coefficients α_3 and β_3 either have opposing signs for the implementation event or are both equal to zero. As the implementation of the MSCI redefinition was undertaken in two steps, the test can be applied to both events. Section 4.3 reports the corresponding empirical results.

An important issue in international finance is the degree of integration of different national stock markets. Are asset prices determined locally or globally (Karolyi and Stulz, 2003)? Frequently, market integration is reviewed indirectly by scrutinizing cross-market ownership. But the prevalent home bias may or may not come with market integration in the asset pricing dimension. This paper directly examines the pricing implications for premium changes and arbitrage risk. Under the hypothesis of national market segmentation, the N assets may be partitioned into M national stock markets. Arbitrage may occur primarily within the national market if the arbitrageurs face trading restrictions with respect to foreign assets. It is straightforward to distinguish the global covariance matrix Σ^G accounting for the full correlation structure between all stocks from a restricted matrix Σ^L which ignores cross-country correlations between stocks in different countries by setting those to zero. Formally, the restricted (local) covariance matrix is defined as

$$(\Sigma^L)_{ij} = \begin{cases} (\Sigma^G)_{ij} & \text{if stocks } i \text{ and } j \text{ are listed in the same country} \\ 0 & \text{otherwise,} \end{cases}$$

where Σ^G denotes the full covariance of all index stock returns. The corresponding local market equity premium change in stock j follows as $[\Sigma^L(w^n - w^o)]_j$ and arbitrage risk as $[\Sigma^L\Sigma^L(w^n - w^o)]_j$. This implies a simple test of international market integration summarized in Proposition 4:

Proposition 4: Integrated versus Segmented Equity Markets
*Let Σ^G denote the global covariance matrix of all asset returns and Σ^L the corresponding covariance matrix with zeros for all cross-country elements. Define incremental (or international) matrices as $\Sigma^{Int} = \Sigma^G - \Sigma^L$ and $\Sigma\Sigma^{Int} = \Sigma^G\Sigma^G - \Sigma^L\Sigma^L$, respectively. The **pre-announcement return** can be decomposed into its local and international components as*

$$\Delta p_1 \approx \alpha_1^L \times \Sigma^L(w^n - w^o) + \alpha_1^{Int} \times \Sigma^{Int}(w^n - w^o) + \beta_1^L \times \Sigma^L\Sigma^L(w^n - w^o) + \beta_1^{Int} \times \Sigma\Sigma^{Int}(w^n - w^o)$$

*and the **post-announcement return*** as

$$\Delta p_2 \approx \beta_2^L \times \Sigma^L\Sigma^L(w^n - w^o) + \beta_2^{Int} \times \Sigma\Sigma^{Int}(w^n - w^o)$$

with

$$\begin{array}{ll} (i) & \alpha_1^L = \alpha_1^{Int} > 0 \quad and \quad \beta_1^L = \beta_1^{Int} < 0 \quad and \quad \beta_2^L = \beta_2^{Int} > 0 \\ (ii) & \alpha_1^L > \alpha_1^{Int} = 0 \quad and \quad \beta_1^L < \beta_1^{Int} = 0 \quad and \quad \beta_2^L > \beta_2^{Int} = 0 \end{array}$$

for (i) complete market integration and (ii) for complete market segmentation, respectively.

Proof: Follows from Propositions 1 and 2 by decomposition of $\Sigma^G$and $\Sigma^G\Sigma^G$.

The intuition behind the test of market integration is straightforward. Assume the stock price of 'Ford' (stock j) is equally strongly correlated with both the stock price of 'GM' (stock g) and the Italian company 'Fiat' (stock f) and that both GM and Fiat are up-weighted in the MSCI index by the same amount, hence $u_g = u_f > 0$. Under market integration, the index weight increase of both GM and Fiat should produce quantitatively the same pre-announcement effect on the stock price of Ford as $\Sigma_{jg}^G u_g = \Sigma_{jf}^G u_f$. This equality of the cross-border pricing effects is tested by separating the GM element $\Sigma_{jg}^G u_g$ as part of the local premium change $\Sigma_{j\bullet}^L u$ from the Fiat element $\Sigma_{jf}^G u_f$ as part of the international premium change $\Sigma_{j\bullet}^{Int} u$. The corresponding regression coefficients are equal ($\alpha_1^L = \alpha_1^{Int}$) if stocks are priced relative to their risk contribution to the global market risk. However, if the risk contribution of Fiat is not part of the market benchmark for the Ford risk premium, then its change should be without consequence for the Ford stock price; hence $\alpha_1^{Int} = 0$. A similar logic applies to the coefficients β_1^L and β_2^L, but with respect to the arbitrageurs. Assume that U.S. stocks are exclusively arbitraged by U.S. investors, Italian stocks by Italian investors, etc.. In this case the sub-matrix $\Sigma^L\Sigma^L$ is sufficient to characterize all arbitrage risk and therefore $\beta^{Int} = 0$. However, the complementary matrix $\Sigma\Sigma^{Int}$ should feature the same price impact ($\beta_1^L = \beta_1^{Int} < 0$ and $\beta_2^L = \beta_2^{Int} > 0$) if arbitrageurs adopt a global arbitrage strategy and treat foreign and home stocks in a similar way. In the latter case, stock markets are integrated with respect to arbitrage behavior.

The above regression specification only explores the average degree of market integration or segmentation. Alternatively, the matrices Σ^{Int} and $\Sigma\Sigma^{Int}$ could be further decomposed into an incremental contribution of each market with respect to all other markets. This allows, in principle, for more specific tests of integration of any particular country either with respect to the world equity market or any other country market. The largest sample and therefore the greatest statistical power is obtained by pooling all observations. The regression results concerning international market integration are reported in Section 4.4.

3 The MSCI Index Redefinition

Morgan Stanley Capital International Inc. (MSCI) is a leading provider of equity (international and U.S.), fixed income and hedge fund indices. The MSCI equity indices are designed to be used by a wide variety of global institutional market participants. They are available in local currency and U.S. Dollars (US$), and with or without dividends reinvested.[10] MSCI's global equity indices have become the most widely used international equity benchmarks by institutional investors. By the year 2000, close to 2,000 organizations worldwide were using the MSCI international equity benchmarks. Over US$ 3 trillion of investments were benchmarked against these indices worldwide and approximately US$ 300 to 350 billion were directly indexed.[11] The index with the largest international coverage is the MSCI ACWI (All Country World Index), which includes

[10] Aggregating individual securities by different criteria MSCI creates a broad base of indexes such as Global, Regional and Country Equity Indexes, Sector, Industry Group and Industry Indexes, Value and Growth Indexes, Small Cap Equity Indexes, Hedged and GDP-weighted Indexes, Custom Equity Indexes, Real Time Equity Indexes.

[11] See the investment newsletter 'Spotlight on: Throwing Weights Around', Hewitt Investment Group, December 2000.

50 developed and emerging equity markets. This broad index is the focus of the empirical work. MSCI reviews the index composition at regular intervals in order to maintain a broad and fair market representation.[12] But in 2000 MSCI initiated a particular index review of exceptional scope described in the following section.

3.1 Announcement of the Index Change

In February 2000, MSCI communicated that it was reviewing its weighting policy and that it was considering a move to index weights defined by the freely floating proportion of the stock value. Such free-float weights would better reflect the limited investibility of many stocks. Free-float weights were consecutively adopted by MSCI's competitor Dow Jones on September 18, 2000. The next day, MSCI published a consultative paper on possible changes and elicited comments from its clients. The consultation process between MSCI and thr investment industry proceeded throughout November 2000. It is therefore very likely that speculators anticipated the change in the index methodology and acquired arbitrage positions prior to the public announcement of the index revision.

This public announcement occurred in two steps. On December 1, MSCI announced that it would communicate its decision on the redefinition of the MSCI international equity index on December 10, 2000. Fund managers could by then infer that MSCI's adoption of free floats weights was extremely likely. The second announcement on December 10, 2000 provided the timetable for the implementation of the index change in two steps and the new target for the market representation of 85 percent up from previously 60 percent. The equity indices would adjust 50 percent towards the new index on November 30, 2001 and the remaining adjustment was scheduled for May 31, 2002. MSCI's decision was broadly in line with the previous consultative paper. Only the target level of 85 percent was somewhat higher (by 5 percent) and the implementation timetable was somewhat longer than most observers had expected.

Investment newsletters suggest that the formal announcement of the index change on December 10, 2000, was largely anticipated by the market.[13] The first announcement on December

[12] The index maintenance can be described by three types of reviews. First, there are annual full country index reviews (at the end of May) in which MSCI re-assesses systematically the various dimensions of the equity universe for all countries. Second, there are quarterly index reviews (at the end of February, August, November), in which other significant market events are accounted for (e.g. large market transactions affecting strategic shareholders, exercise of options, share repurchases, etc.). Third, ongoing event-related changes like mergers and acquisitions, bankruptcies or spin-offs are implemented as they occur.

[13] See again the investment newsletter 'Spotlight on: Throwing Weights Around', Hewitt Investment Group,

1, 2000 provided a strong signal that MSCI had already decided in favor of the free float adoption. The consecutive empirical work explores if pre-announcement arbitrage occurred during the month of November in parallel to MSCI's consultation process. Since the exact beginning of such arbitrage activity is difficult to date, a variety of different event windows are proposed all of which extend until December 1, the date of the first formal announcement. These pre-announcement windows cover alternatively a period of 5, 10, 15 or 20 trading days. After the week-end of December 2 and 3, 2000, the financial market re-opened on December 4. Knowledge of the imminent index change apparently became widespread after the first announcement. Accordingly, the post-announcement event windows starts on December 4 and extends over the following 3, 5 or 7 trading days during which the liquidity providers learn about the index revision. Again, results for alternative window sizes are reported to demonstrate robustness to the exact window length.

It is instructive to plot the return behavior of selected portfolios around December 1, 2000. This evidence can support the particular window choice prior to further cross-sectional analysis. In the light of the theoretical considerations in the previous section, four portfolios are of particular interest. All are constructed as 'self-financing' portfolios with a zero sum of weights and a sum of absolute weights normed to 1:

- The *linear portfolio* has portfolio weights ϖ^{Lin} proportional to the weight change $w^n - w^o$.
- The *premium portfolio* has portfolio weights ϖ^{Pm} proportional to the premium change $\Sigma(w^n - w^o)$.
- The *hedge portfolio* has portfolio weights ϖ^{Hed} proportional to the arbitrage risk $\Sigma\Sigma(w^n - w^o)$.
- The *optimal portfolio* has portfolio weights ϖ^{Opt} proportional to $\Sigma(w^n - w^o) - \theta\Sigma\Sigma(w^n - w^o)$ with $\theta = 0.001$.[14]

Figure 1 graphs these four portfolios for the 7 week period from November 6 to December 22, 2000. A strong positive price reaction is visible after December 4 for the linear portfolio, the premium portfolio, and particularly the hedge portfolio. Noticeable also is the value decline of the hedge portfolio by more than 5 percent from November 6 to December 1, 2000. Short-selling of the portfolio ϖ^{Hed} by arbitrageurs throughout November 2000 can explain this relative

December 2000.

[14] The multi-variable regression analysis in section 4.1. shows that a plausible hedge paramter value is given by $\theta = \beta/\alpha = 0.001$. The latter value is suggested by the parameter estimates for α and β in Table 3.

price decline. The arbitrage model also predicts that the price decline of the hedge portfolio should be reversed once the liquidity suppliers learn about the index change. After December 1, a temporary price increase is found for the hedge portfolio in accordance with proposition 2. Figure 1 also highlights that the linear portfolio shows no price increase during the month of November, which suggest that the speculators did not follow a 'naive' arbitrage strategy based on simple stock weight changes.

Interpreting the pre-announcement return of the premium and hedge portfolio itself is problematic. Proposition 1 predicts a positive pre-announcement return for the premium portfolio and a negative pre-announcement returns for the hedge portfolio *conditional* on controlling for the other portfolio. Only a multi-variable regression analysis can reveal if the pre-announcement return behavior corresponds to proposition 1. But short of a formal regression analysis, I can still examine the return behavior of the optimal portfolio. Returns of the optimal portfolio combine the premium and hedge portfolio and inference is not tainted by a lack of conditioning. The optimal portfolio indeed shows the price run-up predicted in proposition 1. The optimal portfolio increases in value by 6 percent from November 6 to December 1, 2000. Price pressure from the arbitrageurs' position build-up in the optimal portfolio provides an explanation for the steady price increase.

Figures 2 and 3 document the performance of the same 4 portfolios around the first and second implementation day, respectively. There is no evidence for any strong price reaction on either of the two implementation dates. But for the first implementation event, a 2 percent value decrease is noticeable for the 'optimal portfolio' after December 1, 2001. This depreciation of the optimal portfolio could result from an over-estimation of the magnitude of the demand shock ($k > 1$) as stated in proposition 3. The second implementation event is marked by the opposite appreciation of the optimal portfolio. The latter observation could correspond to an underestimation of the magnitude of the demand shock $w^n - w^o$ for the second part of the index adjustment. This evidence is examined again in section 4.3 using cross-sectional regression analysis.

3.2 Overview of the Index Weight Changes

MSCI's new index methodology differs from the previous equity index definition in two aspects. First, stock selection is based on freely floating capital as opposed to market capitalization. Second, the market representation is enhanced in the new index. MSCI defines the free float of a security as the proportion of shares outstanding that is available for purchase by international investors. In practice, limitations on the investment opportunities of international institutions are common due to so-called "strategic holdings" by either public or private investors. Given that disclosure requirements generally do not permit a clear identification of "strategic" investments, MSCI labels shareholdings by classifying investors as strategic and non-strategic. Free floating shares include those held by households, investment funds, mutual funds and unit trusts, pension funds, insurance companies, social security funds and security brokers. The non-free floating shares include those held by governments, companies, banks (excl. trusts), principal officers, board members and employees. The second goal of the equity index modification was an enhanced market representation. In its new indices, MSCI targets a free float-adjusted market representation of 85 percent within each industry and country, compared to the 60 percent share based on market capitalization in the old index. Because of differences in industry structure, the 85 percent threshold may not be uniformly achieved. Moreover, the occasional over- and under-representation of industries may also imply that the aggregate country representation may deviate from the 85 percent target.[15]

Next, I describe the effect of the new index methodology on the index composition. Prior to its revision, the MSCI ACWI included a total of 2077 stocks. The new index methodology led to the inclusion of 489 new stocks and the deletion of 298 stocks. The total number of stocks belonging either to the old or new index is therefore 2566. Table 1 provides a breakdown of these stocks by country and lists the number of retained sample stocks for each country. The sample excludes 62 stocks from the two crisis countries Argentina and Turkey. The analysis also requires 2 years of historic price data to compute covariance matrices with all other index stocks. For 31 stock codes no company information was found. Another 182 stocks have an incomplete price history prior to the index change.[16] This reduces the data sample from 2566 to 2291 stocks, of which 396 are included and 265 excluded in the index revision.

Table 1, columns (3) and (4) provide the aggregate country weight defined as the sum of all stock weights before and after the index revision, respectively. The largest contribution to the new MSCI index comes from the U.S. stocks with 55.12 percent followed by the U.K. with 10.33 percent and Japan with 9.38 percent. The most dramatic country weight change concerns the U.S. with a 6.24 percent absolute weight increase followed by the U.K. with a 1.07 percent

[15] MSCI's bottom-up approach to index construction may lead to a large company in an industry not being included in the index, while a smaller company from a different industry might be included.

[16] I require in particular 80 weekly return observations for the two-year period between July 1, 1998, and July 1, 2000. Otherwise, the return history is incomplete.

increase. Both countries also feature the largest number of new stocks added to the index. Of the 396 sample stocks added to the new MSCI index, a total of 113 are U.S. stocks and 29 are U.K. stocks. It is also instructive to express stock weight changes in percentage terms (relative to the midpoint) as

$$\Delta v_j = \frac{w_j^n - w_j^o}{\frac{1}{2}\left(w_j^n + w_j^o\right)},$$

where w_j^o and w_j^n represent the old and new index weight of stock j, respectively. The percentage weight change is bounded above by 2 for newly included stocks and below by -2 for deleted stocks. Table 1, columns (5) and (6) report the mean and the standard deviation of the percentage weight change Δv_j by country. The largest average stock weight increase is experienced by stocks in New Zealand (44.1 percent), the U.S. (39.0 percent) and the U.K. (36.9 percent). Figure 4 plots the percentage weight change of individual stocks against their initial weight (in logs) both for non-U.S. stocks and U.S. stocks. Due to the overall increase in the number of stocks in the new index, many previously included stocks are down-weighted. This explains why the median percentage weight change is negative at -19.0 percent. The comparison between U.S. and non-U.S. stocks also reveals that the average size of U.S. stocks is larger than for non-U.S. stocks. This size difference applies equally to the groups of added, deleted and re-weighted stocks.

3.3 Risk Premium Changes and Marginal Arbitrage Risk

In order to determine the premium change and the marginal arbitrage risk the covariance matrix Σ of all stock returns needs to be estimated. To proxy for the (expected) covariance matrix, I simply use the historical covariance based on 2 years of return data prior to the event. The estimation window for the covariance covers the period July 1, 1998 to July 1, 2000. It is sufficiently removed from the first announcement on December 1, 2000 to not be affected by the event itself. The covariance estimation for the stock returns is based on weekly data. Since stock prices are sampled around the world, daily sampling may pose inference problems due to asynchronous return measurement. Weekly return sampling appears more robust to this problem and justifies the use of weekly data.[17] On a more general level, using historical data represents certainly an imperfect measure of the forward look covariance, but it is also the mostly likely technique used by arbitrageurs to determine the optimal arbitrage strategy and the ex ante risk

[17] I verify that estimation of the equity return covariance based on a daily return sampling did not qualitatively alter the results.

of their portfolio position. It is important to highlight that even though the covariance matrix is estimated, only a weighted average of its row elements is used to infer the premium change. Every row element $[\Sigma(w^n - w^o)]_j$ is calculated based on approximately 100 weekly observations from 2291 different return sequences. The estimation quality is therefore comparable to the standard beta estimation.

A particularly interesting aspect of the MSCI index revision is its international dimension. The global index change can be interpreted as a natural experiment on local versus global asset pricing. The degree of market integration versus segmentation can be measured in two dimensions. First, I explore whether the cross-sectional price changes around the announcement event correspond to premium changes induced by either local or global beta changes. The international nature of the weight changes assures that local and global beta changes are generally different. If local beta changes alone explain the price behavior, market segmentation is inferred. If the incremental premium changes between the global and the local premium have additional and equal explanatory power, I conclude that global asset pricing and therefore market integration represents the correct benchmark. Second, a similar argument applies to marginal arbitrage risk. If arbitrage strategies are confined to exploiting local premium changes for local stocks, only the risk contribution of local stocks matters. By contrast, global arbitrage strategies optimize over the marginal risk contribution of all local and all international stocks. The marginal risk contribution of local stocks to the portfolio should equal that of the international stocks and both factors should reveal an equal price impact. The two polar cases of market integration and segmentation can be summarized as follows:

1. Global asset pricing and global equity arbitrage: Arbitrageurs take speculative positions in all stocks affected by the index and risk is measured by the global covariance Σ^G of dollar returns. The change in the risk premium on stock j is proportional to $[\Sigma^G(w^n - w^o)]_j$ and the arbitrage risk proxied by $[\Sigma^G\Sigma^G(w^n - w^o)]_j$.

2. Local asset pricing and local equity arbitrage: Arbitrageurs speculate only on the weight change in one local market. I can therefore define a restricted covariance matrix Σ^L of equity returns which is obtained from Σ^G by setting to zero all cross-country covariances. The change in the risk premium under complete market segmentation is proportional to $[\Sigma^L(w^n - w^o)]_j$ and the arbitrage risk proxied by $[\Sigma^L\Sigma^L(w^n - w^o)]_j$.

Table 2 reports summary statistics of the risk premium changes and the corresponding arbitrage risk for different groups of stocks. Panels A and B describe the global and local risk premium change, respectively, while Panels C and D provide summary statistics on global and local arbitrage risk. To better interpret these statistics, the premium change can be related to the corresponding beta change according to

$$\Delta beta_j = -\frac{[\Sigma(w^n - w^o)]_j}{\sigma_m^2}.$$

The weekly market volatility of the global index is estimated as $\sigma_m^2 = w^{o\prime}\Sigma w^o = 0.936$. The global premium change for a stock therefore corresponds to a beta change of similar magnitude and this highlights the large cross-sectional dispersion of beta changes. The standard deviation for the global premium change in Table 2 is given by 0.049, which implies 0.052 for the standard deviation of global beta changes. The MSCI index revision generated a substantial beta change for a large cross-section of stocks.

A graphical representation of the distribution of the global and local risk premium change is provided in Figure 5. It reveals systematic differences between non-U.S. and U.S. stocks. For non-U.S. stocks, the dispersion of the local equity premium change is relatively small and the mean change negative at -0.005. The corresponding average change in the global premium is also negative at -0.009, but features a much higher standard deviation of 0.036 compared to only 0.009 for the local premium. Non-U.S. stocks include more down-weighted than up-weighted stocks, which explains the negative mean for both local and global premium changes. Compared to the global covariance matrix Σ^G, the local covariance matrix Σ^L features by construction many zero elements, which tends to generate less dispersion in the local relative to the global premium change. The dispersion of local premium changes is particularly small for stocks from countries with a minor representation in the MSCI index. It is interesting to note the low correlation between local and global premium changes for non-U.S. stocks. The correlation of local and global premium changes correspond to the correlation of the local and global beta changes and can be calculated as

$$Corr_{j \notin US}\left[\Delta beta_j^L, \Delta beta_j^G\right] = Corr_{j \notin US}\left[\Sigma_{j\bullet}^L(w^n - w^o), \Sigma_{j\bullet}^G(w^n - w^o)\right] = 0.149.$$

This low correlation allows for sufficient discriminatory power between local and global asset pricing effects. Weight changes by other international stocks exercise an important influence on the global beta change for most non-U.S. stocks and therefore differentiate global beta from local beta changes. This aspect underlines that the degree of market integration is very important for the price effect of a global demand shock like the MSCI index revision. Market integration generally implies a completely different price effect for non-U.S. stocks compared to market segmentation.

For U.S. stocks the premium changes behave very differently. The local equity premium change for U.S. stocks shows a positive mean of 0.115 and a large standard deviation of 0.074. The local premium change here is typically only slightly smaller than the global premium change as illustrated in Figure 5. Most U.S. stocks are situated just below the 45 degree line. The large number of U.S. stocks in the MSCI index explain why for U.S. stocks the corresponding rows in the global and local covariance matrices differ less than for stocks from other countries because fewer cross-country covariances are set to zero. As a consequence, local and global premium and beta changes are highly correlated for U.S. stocks; that is

$$Corr_{j\in US}\left[\Delta beta_j^L, \Delta beta_j^G\right] = Corr_{j\in US}\left[\Sigma_{j\bullet}^L(w^n - w^o), \Sigma_{j\bullet}^G(w^n - w^o)\right] = 0.911.$$

This high correlation makes the U.S. stocks less suited for inference about global versus local asset pricing. Intuitively, most of the change in the beta for U.S. stocks is induced by the index weight changes of other U.S. stocks with similar effects on both the local and global betas.

Figure 6 plots the marginal arbitrage risk contribution of each stock under global arbitrage against the marginal risk contribution under local arbitrage. Local arbitrage risk accounts only for the risk of positions in pursuit of local beta changes, whereas global arbitrage risk is related to positions exploiting global beta changes. The distribution of local and global marginal arbitrage risk are closely related to the distribution of the local and global risk premium changes. The marginal arbitrage risk $[\Sigma\Sigma(w^n - w^o)]_j$ differs from the risk premium change only by a quadratic term $\Sigma\Sigma$ replacing the linear term Σ. Again, non-U.S. stocks are found to behave very differently from U.S. stocks. Local and global marginal arbitrage risk have a low correlation of only 0.172 across non-U.S. stocks. However, for U.S. stocks, this correlation is approximately 0.987 and indicates strong colinearity. Meaningful inference about global versus local arbitrage risk is therefore restricted to the sample of non-U.S. stocks.

3.4 Portfolio Risk Relative to the MSCI Index

How much overall risk do arbitrageurs take in their pursuit of an optimal arbitrage strategy? The estimation of the covariance matrices allows not only to assess premium changes and marginal arbitrage risk for each individual stock, but also the portfolio risk involved in the entire arbitrage strategy. It is assumed here for simplicity that the initial holdings of an arbitrageur correspond to the stock weights w^o of the old MSCI index. The risk of such a portfolio can be measured as the standard deviation of the portfolio returns, namely

$$Risk^o = (w^{o\prime}\Sigma w^o)^{\frac{1}{2}}.$$

Next, two alternative portfolios are considered. The first portfolio combines the old index weights with the 'self-financing' linear portfolio ϖ^{Lin} composed of weights proportional to the index weight change $w^n - w^o$. The second portfolio combines the original MSCI weights with the optimal portfolio ϖ^{Opt} which has weights proportional to $\Sigma(w^n - w^o) - \theta\Sigma\Sigma(w^n - w^o)$. A shift from the original portfolio w^o into one of the two arbitrage portfolios implies new portfolio weights given by

$$w^{\kappa Lin} = w^o + \kappa\varpi^{Lin} \quad \text{or} \quad w^{\kappa Opt} = w^o + \kappa\varpi^{Opt},$$

where the parameter κ denotes the leverage factor. For the linear arbitrage portfolio, the factor $\kappa = 1$ corresponds to the full shift into the new index, while $\kappa > 1$ implies a speculative position beyond the weight change. The weight change of the optimal arbitrage portfolio accounts for the changes in the risk premium and the arbitrage risk contribution of each stock. The latter is scaled by $\theta = \beta/\alpha$. For an initial portfolio position w^o, the percentage risk increase due to linear or optimal arbitrage positions with leverage factor κ follows as

$$\frac{Risk^{\kappa Lin}}{Risk^o} = \left[\frac{w^{\kappa Lin\prime}\Sigma w^{\kappa Lin}}{w^{o\prime}\Sigma w^o}\right]^{\frac{1}{2}} \quad \text{or} \quad \frac{Risk^{\kappa Opt}}{Risk^o} = \left[\frac{w^{\kappa Opt\prime}\Sigma w^{\kappa Opt}}{w^{o\prime}\Sigma w^o}\right]^{\frac{1}{2}},$$

respectively. Figure 7 plots the change in the risk ratio as a function of the leverage factor κ for the global arbitrage strategy. A linear portfolio weight shift $w^{\kappa Lin}$ into the new index slightly increases the portfolio risk. This is not surprising since the new index increases the weight of U.S. stocks and their relatively higher correlation diminishes the overall international diversification benefits of the new global equity allocation. The risk increases further with a leverage beyond $\kappa = 1$. For a leverage factor of $\kappa = 3$, the risk of the global arbitrage portfolio increases by 13 percent relative to the old MSCI weights. By contrast, the optimizing arbitrage strategy which shifts into weights $w^{\kappa Opt}$ achieves a substantial risk reduction even as the leverage increases. Based on parameter estimates for α and β, I calibrate $\theta = 0.001$. A risk reduction of 33 percent is found for a leverage factor $\kappa = 3$. This risk reducing effect of the optimal arbitrage strategy compared to a linear arbitrage strategy is due to the term $-\theta\Sigma\Sigma(w^n - w^o)$, which down-weights (up-weights) stocks with a positive (negative) marginal arbitrage risk contributions. The overall portfolio risk of the leveraged optimal arbitrage portfolio is lower than the benchmark portfolio for a wide range of leverage factors κ.

This aspect allows me to clarify a frequent misunderstanding about the limits of arbitrage. The long-term nature of an arbitrage strategy does not necessarily imply that the arbitrage portfolio is excessively risky. In the case of multi-asset demand shocks like the MSCI index revision, the optimizing arbitrage strategy offers an opportunity to reduce absolute portfolio risk exposure relative to the market risk. This absolute risk reduction increases in the time horizon of the arbitrage strategy. Moreover, a price elastic liquidity supply means that expected arbitrage returns are reduced if many arbitrageurs pursue the same arbitrage strategy. However, it will not reduce the risk reduction benefits of the hedge portfolio. A high Sharpe ratio for the arbitrage strategy may therefore come not so much from a high expected excess return, but rather from the reduction of portfolio risk. Focusing only on the (expected or ex-post) excess returns of an arbitrage portfolio may therefore represent a misleading metric for its success.

4 Evidence

The portfolio approach to limited arbitrage allowed me to derive a sequence of testable implications. Event returns in each stock are determined by the change in a stock's risk premium and by its marginal risk contribution to the arbitrage portfolio. Intuitively, the risk premium change $\Sigma(w^n - w^o)$ is given by the product of asset demand change $w^n - w^o$ and the covariance matrix Σ of all stocks. Its j-th element represents the change in the marginal contribution of security j to the total market risk induced by the weight change of all stocks. An arbitrage position proportional to the expected premium change generates total arbitrage risk given by $(w^n - w^o)'\Sigma\Sigma\Sigma(w^n - w^o)$. The marginal contribution of each stock to this arbitrage risk is characterized by $\Sigma\Sigma(w^n - w^o)$. The optimal arbitrage strategy can be proxied as a linear combination of stock weights determined positively by the expected risk premium change $k\Sigma(w^n - w^o)$ and negatively by the expected marginal arbitrage risk $k\Sigma\Sigma(w^n - w^o)$. The optimal arbitrage strategy in combination with the linear liquidity supply trigger proportional pre-announcement event returns as stated in Proposition 1.

Section 4.1 provides the evidence on the pre-announcement effect. Further revisions of the cross-sectional returns occur when the liquidity suppliers learn about the demand shock. These post-announcement return patterns are examined in section 4.2. Upon the implementation of the index change, all market participants experience the true magnitude of the demand shock. Given an under- or overestimation of the shock size, this can give rise to further cross-sectional return patterns documented in section 4.3. The empirical counterpart to Proposition 4 about international market segmentation versus integration is provided in Section 4.4.

4.1 Pre-Announcement Event

The global scale of the MSCI index rebalancing provides an extremely large sample of stocks which experienced a weight change. The sample contains 2291 stocks with a continuous two year price history needed to calculate the global covariance matrix Σ^G. The statistical inference is based on a cross-sectional analysis in which dollar returns Δp_j (defined as log price differences $\ln P_t^j - \ln P_{t-1}^j$) in stock j over the entire event window are regressed on a constant c, the stock's risk premium change $\left[\Sigma^G(w^n - w^o)\right]_j$ and its corresponding marginal arbitrage risk $\left[\Sigma^G\Sigma^G(w^n - w^o)\right]_j$. Formally,

$$\Delta p_1^j = c + \alpha_1 \times \left[\Sigma^G(w^n - w^o)\right]_j + \beta_1 \times \left[\Sigma^G\Sigma^G(w^n - w^o)\right]_j + \mu_j,$$

where clustering of the error term μ_j on the country level is allowed. Error clustering at the country level can account for omitted exchange rate effects or common country effects.

It is difficult to know when arbitrage trading on the index revision started. Four alternative pre-announcement windows are considered comprising 5, 10, 15 or 20 trading days prior to December 1, 2000. Table 3, Panel A, features the regression results for the full sample of 2291 stocks. Reported are regression results with a specification including only the constant and the risk premium change as well as the complete specification. A specification without the marginal arbitrage risk term corresponds to the nested Greenwood model. This specification is correct for the special case $\lambda = 1$ where all market participants are equally informed arbitrageurs and there is no liquidity supply. The restrictive specification is strongly rejected by the data. The estimated coefficient α_1 is negative while theory predicts a positive coefficient. The rejection of the Greenwood model is evident for each of the four pre-announcement event windows. But under the full specification with the arbitrage risk term, the sign of the coefficient α_1 becomes positive at a high level of statistical significance. The coefficient estimate of 80.6 for the 10 day event window also implies an economically large return difference of approximately 3.95 percent for two stocks with a relative change in their risk premium by one standard deviation or 0.049. The coefficient β_1 also takes on the predicted negative sign with a value of -0.099 for the 10 day event window. This means that an arbitrage risk increase by one standard deviation (or 61.63) in a particular stock induces smaller speculative positions and therefore a decrease in the 10 day pre-announcement return by 6.1 percent. The adjusted R-squared of the full specification is at 0.119 highest for the 20 day event window and more than 2 times higher than under the restrictive specification. The estimated coefficients for the full specification increase in the window size as can be expected if the return effects of arbitrage cumulate over time.

As a robustness check, Panel B reports the results for the sample of added and deleted stocks and Panel C for non-U.S. stocks only. Both samples feature qualitatively similar results. In each case and for every window size, the hypothesis that $\beta_1 = 0$ is strongly rejected. As in the entire sample, and in line with the theoretical model, the coefficient α_1 for the risk premium change is significantly positive and the coefficient β_1 for the arbitrage risk significantly negative in the full specification. In Panel B the adjusted R-squared is generally higher, suggesting a better model fit for stocks with the most dramatic weight changes. For the 20 day window in Panel B, an adjusted R-squared of 0.18 is found. Overall, pre-announcement returns provide strong empirical support for the generalized arbitrage model. The estimated effects are also economically significant.

4.2 Post-Announcement Event

Proposition 2 asserts that the return effect of the hedge position risk is reversed once the liquidity suppliers learn about the index revision. Arbitrageurs then liquidate their arbitrage positions. It is assumed that knowledge of the index revision became widespread after the first MSCI announcement on December 1, 2000. Three alternative post-announcement event windows are defined to calculate the cumulative returns. These event windows all start on December 4, 2000 (after the week-end of December 2 and 3) and cover return over 3, 5 or 7 trading days. From Proposition 2, the preferred cross-sectional specification for the post-announcement return effect follows as

$$\Delta p_2^j = c + \beta_2 \times \left[\Sigma^G \Sigma^G (w^n - w^o)\right]_j + \mu_j,$$

where $\beta_2 > 0$ is expected for the general model with $\lambda < 0$. The cross-sectional price effect is generated by the liquidation of hedging positions. The nested Greenwood model which abstracts from speculative position taking and hedging implies $\beta_2 = 0$ under $\lambda = 1$.

Table 4, Panel A, reports regression results for the base specification and an augmented specification which includes the premium change $\Sigma^G(w^n - w^o)$ as an additional control variable. The reported t-statistics are robust to error clustering at the country level. The coefficient β_2 is significant at the 1 percent level in all specifications, all samples and for all three event windows. The adjusted R-squared for the base specification reaches 0.138 for the 5 day event window. It is at 0.175 even higher for the sample of added and deleted stocks reported in Panel B. The premium effect captured by the coefficient α_2 is significant on a 1 percent level for the full sample, but not for the sample of added and deleted stocks or the sample of non-U.S. stocks. The Greenwood model with $\beta_2 = 0$ is again strongly rejected by the data.

The post-announcement return pattern provides additional support for the generalized model of risk arbitrage. It corresponds to the evidence in section 4.1 which shows that the pre-announcement returns are also modified by marginal arbitrage risk. The speculative dynamics around announcement of the index revision are therefore best captured in a model which features uninformed liquidity suppliers. The baseline CAPM or Greenwood framework which ignores information asymmetries cannot account for the observed price pattern. In particular, the price significance of the hedging demand around the announcement event suggests that hedging of the arbitrage risk was an important element of the speculative strategy of the arbitrageurs.

4.3 Implementation Effects due to Demand Shock Uncertainty

While the weight change of individual stocks might have been quite predictable, the same cannot be said about the exact magnitude of the demand shock. The magnitude of the demand shock may be uncertain because the value of all index tracking wealth is relatively hard to predict. Moreover, many funds might have had some discretion over the exact timing of the index revision since the old and the new index coexisted for the period between the first and second implementation date. It is therefore very plausible that the arbitrageurs' beliefs $\widehat{u} = ku = \widetilde{\mathcal{E}}(u)$ about the magnitude of the shock differ from the correct beliefs u by some factor k, where $k > 1$ corresponds to an overestimation of the shock and $0 < k < 1$ to its underestimation. The implementation of the index revision on November 30, 2001, and May 31, 2002, naturally provides new information about u and allows for more precise posterior beliefs. Proposition 3 distinguishes the return effect resulting from prior underestimation and overestimation. Opposite signs for the coefficients α_3 and β_3 are expected in both cases.

Two alternative window lengths are chosen for the two implementation event. Event windows measuring the cumulative daily stock return start two days before the implementation dates of November 30, 2001, and May 31, 2002, respectively, and extend over the next 5 or 7 trading days. The two-step implementation process for the MSCI index revision provides two separate observations to examine predication errors about the size of the demand shock. In Table 4, Panels A and B report the evidence for the first implementation date and Panels C and D for the second implementation date. For both the 5 and 7 day event windows in Panel A, the coefficient α_3 on the risk premium change has a negative sign and is economically large. The coefficient β_3 for the arbitrage risk has the opposite positive sign as predicted for the case of an overestimation of the demand shock magnitude. The results are very similar for the sample of added and deleted stocks (Panel B). For the second implementation event I find (in absolute terms) smaller coefficients with a positive parameter estimate for α_3 and a negative estimate for β_3. This suggests that the second demand shock was underestimated contrary to the first one. Overall, no evidence is found which implies a model rejection since α_3 and β_3 have opposite signs in each regression. The economically and statistically significant coefficient estimates around the implementation events also suggest parameter uncertainty with respect to the magnitude of the demand shock.

4.4 Global versus Local Asset Pricing

Arbitrage strategies could comprise all MSCI stocks or only a subset of re-weighted stocks in the local market. The investor mandate might constrain some fund managers not to invest in the foreign equity market. Similarly, dedicated country funds may be limited to investment in only one foreign country. Only a local equity arbitrage strategy is feasible in these cases. In order to discriminate between the role of local and global asset pricing, the incremental international risk premium change is defined as

$$\left[\Sigma^{Int}(w^n - w^o)\right]_j = \left[\Sigma^G(w^n - w^o)\right]_j - \left[\Sigma^L(w^n - w^o)\right]_j,$$

and the incremental international marginal arbitrage risk as

$$\left[\Sigma\Sigma^{Int}(w^n - w^o)\right]_j = \left[\Sigma^G\Sigma^G(w^n - w^o)\right]_j - \left[\Sigma^L\Sigma^L(w^n - w^o)\right]_j,$$

where Σ^G represents the covariance of dollar returns for all 2291 stocks and Σ^L the equivalent covariance matrix with zeros for stocks in different countries. The statistical inference for the pre-announcement event is based on the regressions

$$\begin{aligned}\Delta p_1^j &= c + \alpha_1^L \times \left[\Sigma^L(w^n - w^o)\right]_j + \alpha_1^{Int} \times \left[\Sigma^{Int}(w^n - w^o)\right]_j + \\ &\quad + \beta_1^L \times \left[\Sigma^L\Sigma^L(w^n - w^o)\right]_j + \beta_1^{Int} \times \left[\Sigma\Sigma^{Int}(w^n - w^o)\right]_j + \mu_j,\end{aligned}$$

and for the post-announcement return on

$$\Delta p_2^j = c + \beta_2^L \times \left[\Sigma^L\Sigma^L(w^n - w^o)\right]_j + \beta_2^{Int} \times \left[\Sigma\Sigma^{Int}(w^n - w^o)\right]_j + \mu_j,$$

where Δp_1^j and Δp_2^j denote the cumulative dollar return for the respective event windows. The coefficient α_1^L measures the return effect of the local premium change and α_1^{Int} the incremental premium change if stocks are priced globally. Similarly, β^L and β^{Int} capture the marginal arbitrage risk effect on returns for the local arbitrageur and the incremental effect for the global arbitrageur, respectively. Equality of the coefficients α_1^L and α_1^{Int} implies global asset pricing and equality of β_1^L and β_1^{Int} (as well as β_2^L and β_2^{Int}) implies global arbitrage. Both suggest an integrated global equity market. However, $\alpha_1^{Int} = 0$ suggests local asset pricing and $\beta^{Int} = 0$ strictly local arbitrage strategies. The latter two cases characterize an internationally segmented market.

Table 6 reports regression results for the decomposition into the local and global pre-announcement return components. In Panel A, the sample consists of all stocks. The incremental effects captured by the coefficients α_1^{Int} and β_1^{Int} are significant for each of the event windows and have the expected sign. The risk premium change and the marginal arbitrage risk therefore have a significant international component. The arbitrage strategies therefore assumed the validity of an international premium change and also engaged in international hedging. The last two columns in Table 7 report the significance level for an F-test conjecturing equality of the respective coefficients. The null hypothesis $\alpha_1^L = \alpha_1^{Int}$ as well as $\beta_1^L = \beta_1^{Int}$ cannot be rejected. For the 15 and 20 day event windows, surprisingly similar coefficient estimates are obtained. The local beta change for example has a coefficient estimate $\alpha_1^L = 144.8$ for the 20 day return in Panel A and the complementary international beta change (induced by weight changes in foreign country stocks) has a coefficient estimate $\alpha_1^{Int} = 130.5$. The corresponding estimates for the marginal arbitrage risk are $\beta_1^L = -0.186$ and $\beta_1^{Int} = -0.154$.

As a robustness check, separate results are estimated for the smaller sample of added and deleted stocks. These results are reported in Panel B and are qualitatively similar. For the long window of 15 and 20 trading days (in absolute terms) larger point estimates for the local coefficients α_1^L and β_1^L are obtained. But the differences to respective international coefficients α_1^{Int} and β_1^{Int} remain statistically insignificant. An alternative sample is formed by all non-U.S. stocks. U.S. stocks are characterized by a relatively high correlation between local and global risk premium changes as well as between local and global marginal arbitrage risk (see Figures 5 and 6). This makes discrimination between the local and global pricing component more difficult. Non-US stocks feature a much lower correlation between local and global explanatory variable. On the other hand, their local premium and local arbitrage risk variation is small and the coefficient $\alpha_1^L$and β_1^L therefore statistically insignificant for all regressions in Panel C. However, the incremental international coefficients α_1^{Int} and β_1^{Int} are of the predicted sign and statistically different from zero for all event windows. Similar to the full sample, the hypothesis of equity market integration cannot be rejected, but the hypothesis of local asset pricing is strongly rejected.

Table 7 reports the corresponding regression on local versus global pricing for the post-announcement period. In the full sample, the coefficient β_2^{Int} is again highly significant with the correct positive sign. Its magnitude is similar to the local arbitrage risk coefficient β_2^L for both the full sample (Panel A) and the sample of added and deleted stocks (Panel B). For the 5 and 7 day window the null hypothesis $\beta_2^L = \beta_2^{Int}$ cannot be rejected. Only the three day window shows a statistically significant difference. But it is the international coefficient which is largest and this cannot be interpreted as evidence for market segmentation. In the sample of non-U.S. firms (Panel C), only the international coefficient is significant. This is not surprising since local marginal arbitrage risk features hardly any cross-sectional variation among non-U.S. stocks. Overall, post-announcement returns provide additional support in favor of market integration.

5 Conclusion

The previous finance literature viewed equity index changes as an interesting exogenous event to explore the limits of equity arbitrage. This literature has produced evidence for important liquidity effects related to demand shocks. But large-scale multi-asset demand shocks also have a more fundamental interpretation as exogenous changes to the stock-specific risk premium. This paper develops the cross sectional asset pricing implications if the demand shock coming from the index investors is partially anticipated and speculative position taking occurs prior to public announcement of the index change. A new and simple heterogenous agent model of multi-asset arbitrage is proposed. Incorporating information heterogeneity between arbitrageurs and the liquidity supply side of the market has attractive theoretical implications. Arbitrageurs can actually accumulate speculative positions against the less informed liquidity providers. As a consequence, their trading returns can exceed the CAPM-based fair risk compensation. This brings theory closer to a practitioner's understanding of arbitrage, but has also interesting theoretical and empirical implications.

The optimal arbitrage strategy in the generalized setting consists of a trade-off between higher expected returns and lower arbitrage risk. To a linear approximation, the optimal portfolio can be represented as a combination of a 'premium seeking portfolio' and a 'hedge portfolio'. Both portfolio components have distinct cross-sectional asset pricing implications. First, a CAPM price effect is obtained for each stock j which is proportional to the (expected) premium change $[\Sigma(w^n - w^o)]_j$. It reflects the changes in the stock's beta and occurs when arbitrageurs anticipate the index revision. Second, the ability of arbitrageurs to control arbitrage risk via a hedge portfolio generates an additional price effect. The optimal arbitrage strategy consists in modifying stock weights according to their marginal arbitrage risk contribution. Intuitively, a premium seeking portfolio $\Sigma(w^n - w^o)$ generates an absolute arbitrage risk $(w^n - w^o)'\Sigma\Sigma\Sigma(w^n - w^o)$ and a marginal arbitrage risk contribution characterized by $\Sigma\Sigma(w^n - w^o)$. Short selling of the 'hedge portfolio' proportional to $\Sigma\Sigma(w^n - w^o)$ represents the optimal risk reduction for the arbitrageur. Selling of the hedge portfolio during the pre-announcement period implies a negative return followed by a positive return in the post-announcement period when the same short positions are liquidated.

The redefinition of the MSCI index represents an ideal experiment to test the generalized portfolio approach to limited arbitrage. The unprecedented scope of the index revision provides a sample of 2291 stocks for which the covariance matrix Σ can be estimated and for which premium changes and marginal arbitrage risk contributions can be calculated. An important finding in this paper is that pre-announcement returns are determined positively by premium change and negatively by the marginal arbitrage risk contribution. Both are statistically and economically significant explanatory variables for the cross section of returns. For the post-announcement period, the marginal arbitrage risk is also found to have the predicted positive effect on the cross-section of event returns. These findings are robust to variations of the event window size and extend to various subsamples.

The international nature of the MSCI index turns its revision also into a test of global versus local asset pricing. The global covariance matrix Σ^G of all stocks can be decomposed into (i) a covariance matrix Σ^L consisting only of covariances of local stocks domiciled in the same national market and (ii) a complementary matrix $\Sigma^{Int} = \Sigma^G - \Sigma^L$ capturing the effect of international market integration. An important finding for MSCI index revision is that local premium changes alone cannot account for cross-section of price changes around the announcement event. The international component of the premium changes $\Sigma^{Int}(w^n - w^o)$ is statistically highly significant and of similar magnitude. This allows us to reject the hypothesis of market segmentation. Asset returns are best captured by global and not local beta changes. Asset pricing models using a global benchmark appear more appropriate than models based on a local market benchmark. A similar conclusion is reached with respect to arbitrage risk. The international component $\Sigma\Sigma^{Int}(w^n - w^o)$ to the marginal arbitrage risk represents a highly significant pricing factor. This suggests that arbitrage strategies for the MSCI revision were implemented globally.

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Appendix

Proposition 1:

The model is solved backwards starting at the terminal asset value p_4. The risk premium for the last period (after the demand shock $u = w^n - w^o$) follows from market clearing at time $t = 3$ as

$$r_4 = \left[\lambda(\rho\Sigma)^{-1} + (1-\lambda)\gamma I\right]^{-1}(S-u) = \left(I + (1-\lambda)\gamma\frac{\rho}{\lambda}\Sigma\right)^{-1}\frac{\rho}{\lambda}\Sigma(S-u),$$

and the price therefore as

$$p_3 = 1 + \varepsilon_1 + \varepsilon_2 + \varepsilon_3 - r_4.$$

At time $t = 2$, both the arbitrageurs and the liquidity suppliers hold expectations $\widetilde{\mathcal{E}}_2(u) = \overline{\mathcal{E}}_2(u) = \widehat{u}$. The risk premium for period 3 is determined by the asset supply S and therefore given by

$$r = \left[\lambda(\rho\Sigma)^{-1} + (1-\lambda)\gamma I\right]^{-1} S = \left(I + (1-\lambda)\gamma\frac{\rho}{\lambda}\Sigma\right)^{-1}\frac{\rho}{\lambda}\Sigma S,$$

and the price follows as

$$p_2 = 1 + \varepsilon_1 + \varepsilon_2 - \widehat{r}_4 - r,$$

where $\widehat{r}_4 = \left[\lambda(\rho\Sigma)^{-1} + (1-\lambda)\gamma I\right]^{-1}(S-\widehat{u})$ denotes the expected risk premium in period 4.

At time $t = 1$, only the arbitrageurs anticipate the demand shock. Asset price expectations therefore differ and follow as $\widetilde{\mathcal{E}}_1(p_2) = 1+\varepsilon_1 - \widehat{r}_4 - r$ for the arbitrageurs and $\overline{\mathcal{E}}_1(p_2) = 1+\varepsilon_1 - 2r$ for the liquidity suppliers. Market clearing then implies

$$\begin{aligned}
\Sigma S &= \lambda\frac{1}{\rho}\widetilde{\mathcal{E}}_1(p_2-p_1) + (1-\lambda)\gamma\Sigma\overline{\mathcal{E}}_1(p_2-p_1) \\
&= \lambda\frac{1}{\rho}I\left(1+\varepsilon_1 - \widehat{r}_4 - r - p_1\right) + (1-\lambda)\gamma\Sigma(1+\varepsilon_1 - 2r - p_1) \\
&= (\lambda\frac{1}{\rho}I + (1-\lambda)\gamma\Sigma)(1+\varepsilon_1 - p_1) + \lambda\frac{1}{\rho}I\left(-\widehat{r}_4 - r\right) + (1-\lambda)\gamma\Sigma(-2r)
\end{aligned}$$

or

$$\begin{aligned}
p_1 &= 1 + \varepsilon_1 + \left(\frac{\lambda}{\rho}I + (1-\lambda)\gamma\Sigma\right)^{-1}\left[\frac{\lambda}{\rho}I\left(-\widehat{r}_4 - r\right) + (1-\lambda)\gamma\Sigma(-2r) - \Sigma S\right] \\
&= 1 + \varepsilon_1 - 3r + \left(\frac{\lambda}{\rho}I + (1-\lambda)\gamma\Sigma\right)^{-2}\frac{\lambda}{\rho}\Sigma\widehat{u} \\
&= 1 + \varepsilon_1 - 3r + (I + \theta\Sigma)^{-2}\frac{\rho}{\lambda}\Sigma\widehat{u},
\end{aligned}$$

where $\theta = (1-\lambda)\gamma\frac{\rho}{\lambda}$. For $(1-\lambda)\gamma \approx 0$, a first order approximation implies

$$(I+\theta\Sigma)(I-\theta\Sigma) = I - \theta^2\Sigma\Sigma \approx I,$$

and multiplying both sides by $(I+\theta\Sigma)^{-1}$ gives $(I-\theta\Sigma) \approx (I+\theta\Sigma)^{-1}$. It follows that

$$(I+\theta\Sigma)^{-2} \approx (I-\theta\Sigma)^2 = I - 2\theta\Sigma + \theta^2\Sigma^2 \approx I - 2\theta\Sigma.$$

In the absence of the liquidity demand shock the price follows as $p_1(u=0) = 1+\varepsilon_1 - 3r$. The pre-announcement price change is then given by

$$\begin{aligned}
\Delta p_1 &= (I-2\theta\Sigma)^{-2}\frac{\rho}{\lambda}\Sigma\widehat{u} \\
&= \frac{\rho}{\lambda}\Sigma\widehat{u} - 2\frac{\rho}{\lambda}\theta\Sigma\Sigma\widehat{u} \\
&= \frac{\rho}{\lambda}k\Sigma(w^n - w^o) - 2(1-\lambda)k\left(\frac{\rho}{\lambda}\right)^2\gamma\Sigma\Sigma(w^n - w^o),
\end{aligned}$$

where $\widehat{u} = ku$. The aggregate position of the arbitrageurs can be expressed as

$$\begin{aligned}
x^A &= S - (1-\lambda)\gamma\overline{\mathcal{E}}_1(p_2 - p_1) \\
&= S - (1-\lambda)\gamma\left[r - \frac{\rho}{\lambda}k\Sigma u + 2(1-\lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2 k\Sigma\Sigma u\right] \\
&= S - (1-\lambda)\gamma r + (1-\lambda)\gamma\frac{\rho}{\lambda}k\Sigma u - 2(1-\lambda)^2\gamma^2\left(\frac{\rho}{\lambda}\right)^2 k\Sigma\Sigma u
\end{aligned}$$

compared to $x^A(\lambda = 1) = S$ for the case where there are no liquidity providers. The term $k\Sigma u$ represents position taking due to expected premium changes and $k\Sigma\Sigma u$ proxies the expected arbitrage risk. The optimal arbitrage position is given by a linear combination of the expected premium change and the arbitrage risk.

Proposition 2:

From proposition 1, the following price dynamics are obtained:

$$\begin{aligned}
p_0 &= 1 - 4r \\
p_1 &\approx 1 + \varepsilon_1 - 3r + \frac{\rho}{\lambda}k\Sigma u - 2(1-\lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2 k\Sigma\Sigma u \\
p_2 &= 1 + \varepsilon_1 + \varepsilon_2 - r - \widehat{r}_4 \\
p_3 &= 1 + \varepsilon_1 + \varepsilon_2 + \varepsilon_3 - r_4.
\end{aligned}$$

The price at time $t = 2$ without demand shock is given by $p_2(u=0) = 1 + \varepsilon_1 + \varepsilon_2 - 2r$. The arbitrage demand changes the stock price by Δp_1 so that the remaining adjustment is given by

$$\begin{aligned}
\Delta p_2 &= p_2 - p_2(u=0) - \Delta p_1 \\
&\approx -\widehat{r}_4 + r - \frac{\rho}{\lambda}\Sigma\widehat{u} + 2\frac{\rho}{\lambda}\theta\Sigma\Sigma\widehat{u} \\
&= \frac{\rho}{\lambda}\theta\Sigma\Sigma\widehat{u} \\
&= (1-\lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2 k\Sigma\Sigma(w^n - w^o),
\end{aligned}$$

using the substitution

$$\begin{aligned}
r - \widehat{r}_4 &= \left[\lambda(\rho\Sigma)^{-1} + (1-\lambda)\gamma I\right]^{-1}\widehat{u} \\
&= \left[I + (1-\lambda)\gamma\frac{\rho}{\lambda}\Sigma\right]^{-1}\frac{\rho}{\lambda}\Sigma\widehat{u} \\
&\approx \frac{\rho}{\lambda}\Sigma\widehat{u} - \frac{\rho}{\lambda}\theta\Sigma\Sigma\widehat{u}.
\end{aligned}$$

Proposition 3:

Immediately before the correct magnitude of the demand shock becomes known at time $t = 3$, the stock prices are given by

$$\begin{aligned}
p_{3-\Delta t} &= 1 + \varepsilon_1 + \varepsilon_2 + \varepsilon_3 - \widehat{r}_4 \\
&= 1 + \varepsilon_1 + \varepsilon_2 + \varepsilon_3 - \left[\lambda(\rho\Sigma)^{-1} + (1-\lambda)\gamma I\right]^{-1}(S - \widehat{u})
\end{aligned}$$

and the implementation effect becomes (under a linear approximation)

$$\begin{aligned}
\Delta p_3 &= -r_4 + \widehat{r}_4 \\
&= \left[\frac{\lambda}{\rho}\Sigma^{-1} + (1-\lambda)\gamma I\right]^{-1}(u - \widehat{u}) \\
&= \left[I + (1-\lambda)\gamma\frac{\rho}{\lambda}\Sigma\right]^{-1}\left(\frac{\rho}{\lambda}\Sigma\right)(1-k)u \\
&\approx \frac{\rho}{\lambda}(1-k)\Sigma(w^n - w^o) - (1-\lambda)\gamma\left(\frac{\rho}{\lambda}\right)^2(1-k)\Sigma\Sigma(w^n - w^o).
\end{aligned}$$

Table 1: Summary Statistics on Weight Changes

Reported are summary statistics by country on the (1) total number of stocks concerned by MSCI index revision, (2) total number of sample stocks with complete historic price data, (3) new and (4) old country weights in percent. For the sample stocks we also provide the (5) mean and (6) standard deviation of the percentage weight change $\Delta v = 2(w^n - w^o)/(w^n + w^o)$ within the country.

Country	(1) MSCI Stocks	(2) Sample Stocks	(3) New Weight	(4) Old Weight	(5) Mean(Δv)	(6) SD(Δv)
Argentina	17	0	0.03	0.07	–	–
Australia	71	66	1.28	1.29	0.395	0.939
Austria	17	16	0.04	0.09	−0.702	1.021
Belgium	18	16	0.28	0.39	−0.308	0.991
Brazil	50	47	0.30	0.49	−0.638	1.041
Canada	89	80	1.93	1.97	0.195	1.025
Chile	30	30	0.07	0.18	−0.777	0.941
China	43	37	0.26	0.35	0.091	1.215
Colombia	9	8	0.00	0.01	−1.184	0.560
Czech Republic	6	6	0.01	0.03	−0.958	0.149
Denmark	27	24	0.28	0.40	0.098	1.211
Egypt	14	11	0.01	0.01	0.033	0.000
Finland	30	27	1.00	0.98	−0.598	1.116
France	63	59	3.77	4.93	−0.279	1.03
Germany	59	51	2.76	3.71	−0.276	1.013
Greece	24	21	0.16	0.16	0.033	0.000
Hong Kong	34	32	0.66	0.93	−0.511	1.137
Hungary	13	13	0.03	0.04	−1.079	0.913
India	76	74	0.12	0.35	−1.037	0.943
Indonesia	32	32	0.02	0.03	−1.171	1.039
Ireland	16	15	0.34	0.30	0.044	0.942
Israel	52	49	0.16	0.24	−0.765	1.407
Italy	47	41	1.38	1.99	−0.436	1.107
Japan	348	333	9.38	10.71	0.035	1.068
Korea	82	69	0.45	0.57	−0.119	1.102
Luxembourg	3	3	0.01	0.05	−1.093	0.478
Malaysia	86	84	0.12	0.29	−0.696	1.186
Mexico	27	25	0.3	0.59	−0.674	0.938
Morocco	12	11	0.01	0.01	0.033	0.000
Netherlands	26	24	2.17	2.42	−0.045	0.665
New Zealand	15	10	0.04	0.05	0.441	1.446
Norway	24	22	0.16	0.20	0.007	0.834
Pakistan	18	18	0.00	0.01	−1.339	0.706
Peru	11	10	0.01	0.02	−0.294	1.344
Philippines	21	20	0.02	0.05	−0.901	1.176
Poland	22	19	0.04	0.06	−0.404	0.945
Portugal	11	10	0.15	0.23	−0.728	0.492
Russia	12	9	0.08	0.14	−0.550	1.178
Singapore	40	34	0.26	0.38	−0.221	1.167
South Africa	48	45	0.47	0.55	−0.378	0.922
Spain	34	31	1.21	1.38	−0.830	0.964
Sri Lanka	8	8	0.00	0.00	−0.753	0.589
Sweden	37	35	0.84	1.00	−0.022	0.647
Switzerland	43	38	2.87	2.93	−0.105	1.143
Taiwan	96	89	0.44	0.69	0.111	1.304
Thailand	39	37	0.04	0.08	−0.926	1.176
Turkey	45	0	0.05	0.12	–	–
United Kingdom	140	133	10.33	9.26	0.369	0.924
United States	443	414	55.12	48.88	0.390	1.108
Venezuela	7	5	0.01	0.02	−0.598	1.470
Country Unknown	31	0	0.5	0.38	–	–
Total	2566	2291	100.00	100.00	–	–

Table 2: Summary Statistics on Premium Changes and Marginal Arbitrage Risk

Reported are summary statistics on stock risk premium changes and on their risk contributions to the arbitrage portfolio for both the global covariance matrix Σ^G and local covariance matrix Σ^L of stock returns. In the local covariance matrix elements are set to zero for stocks in different national markets. The covariance matrices are estimated for 2 years of weekly dollar stock returns for the period of July 1, 1998 to July 1, 2000. The weekly return variance of the global index is estimated as 0.936. Global premium changes therefore correspond approximately to global stock beta changes.

		Obs.	Mean	S.D.	Min	Max
Panel A: Change in Risk Premium under Global Pricing						
All Stocks	$[\Sigma^G(w^n - w^o)]_j$	2,291	0.006	0.049	−0.173	0.249
Added and Deleted Stocks	$[\Sigma^G(w^n - w^o)]_j$	661	0.013	0.057	−0.173	0.280
U.S. Stocks	$[\Sigma^G(w^n - w^o)]_j$	414	0.070	0.047	−0.078	0.249
Non-U.S. Stocks	$[\Sigma^G(w^n - w^o)]_j$	1877	−0.009	0.036	−0.173	0.219
Panel B: Change in Risk Premium under Local Pricing						
All Stocks	$[\Sigma^L(w^n - w^o)]_j$	2,291	0.017	0.056	−0.074	0.345
Added and Deleted Stocks	$[\Sigma^L(w^n - w^o)]_j$	661	0.027	0.072	−0.071	0.345
U.S. Stocks	$[\Sigma^L(w^n - w^o)]_j$	414	0.115	0.074	−0.074	0.345
Non-U.S. Stocks	$[\Sigma^L(w^n - w^o)]_j$	1877	−0.005	0.009	−0.052	0.031
Panel C: Risk Contribution to Global Arbitrage Portfolio						
All Stocks	$[\Sigma^G\Sigma^G(w^n - w^o)]_j$	2,291	25.88	61.63	−179.46	335.31
Added and Deleted Stocks	$[\Sigma^G\Sigma^G(w^n - w^o)]_j$	661	33.36	72.61	−179.46	335.31
U.S. Stocks	$[\Sigma^G\Sigma^G(w^n - w^o)]_j$	414	100.49	66.14	−84.17	335.31
Non-U.S. Stocks	$[\Sigma^G\Sigma^G(w^n - w^o)]_j$	1877	9.42	46.62	−179.46	302.09
Panel D: Risk Contribution to Local Arbitrage Portfolio						
All Stocks	$[\Sigma^L\Sigma^L(w^n - w^o)]_j$	2,291	20.35	57.00	−71.15	391.95
Added and Deleted Stocks	$[\Sigma^L\Sigma^L(w^n - w^o)]_j$	661	30.12	74.09	−71.15	391.95
U.S. Stocks	$[\Sigma^L\Sigma^L(w^n - w^o)]_j$	414	119.30	77.46	−71.15	391.95
Non-U.S. Stocks	$[\Sigma^L\Sigma^L(w^n - w^o)]_j$	1877	−1.47	2.90	−18.69	11.77

Table 3: Pre-Announcement Event

The cumulative pre-announcement equity returns Δp_1^j (denominated in dollars and expressed in percentage points) over different even windows (WS) is regressed on a constant, the change in the risk premium $\left[\Sigma^G(w^n - w^o)\right]_j$ and the arbitrage risk $\left[\Sigma^G\Sigma^G(w^n - w^o)\right]_j$ of each stock j. Formally,

$$\Delta p_1^j = c + \alpha_1 \times [\Sigma^G(w^n - w^o)]_j + \beta_1 \times [\Sigma^G\Sigma^G(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for 2 years of weekly dollar stock returns for the period of July 1, 1998 to July 1, 2000. The event window size is chosen alternatively to start WS = $5, 10, 15, 20$ trading days prior to December 1, 2000. Panel A reports the coefficients for the entire sample, Panel B for only the added and deleted stocks and Panel C for the subsample of non-U.S. stocks. Robust and country clustered adjusted t-values are reported in parenthesis.

WS	c	[t]	α_1	[t]	β_1	[t]	R^2
Panel A: Announcement Event (All Stocks, N=2291)							
5	0.31	[0.59]	−33.7	[−3.77]			0.054
5	1.54	[3.15]	41.8	[3.42]	−0.064	[−6.54]	0.095
10	−2.14	[−2.98]	−36.0	[−3.26]			0.034
10	−0.25	[−0.32]	80.6	[4.01]	−0.099	[−6.43]	0.089
15	−4.06	[−3.97]	−22.5	[−1.97]			0.009
15	−2.04	[−1.58]	101.7	[3.71]	−0.105	[−4.45]	0.051
20	−5.09	[−5.01]	−65.8	[−4.44]			0.052
20	−1.99	[−1.34]	124.5	[3.47]	−0.161	[−4.58]	0.119
Panel B: Announcement Event (Only Added and Deleted Stocks, N=661)							
5	−0.18	[−0.32]	−44.2	[−4.92]			0.098
5	1.10	[2.00]	42.4	[2.45]	−0.072	[−5.10]	0.145
10	−2.61	[−3.39]	−50.7	[−6.09]			0.077
10	−0.87	[−0.85]	67.2	[2.17]	−0.098	[−4.10]	0.129
15	−4.68	[−3.91]	−37.7	[−3.62]			0.029
15	−2.70	[−1.55]	96.4	[2.39]	−0.111	[−3.11]	0.076
20	−5.04	[−4.21]	−95.7	[−6.20]			0.115
20	−2.05	[−1.03]	106.9	[1.80]	−0.168	[−3.12]	0.180
Panel C: Announcement Event (Only Non-U.S. Stocks, N=1877)							
5	0.28	[0.47]	−27.5	[−1.70]			0.022
5	1.43	[2.70]	41.2	[2.71]	−0.060	[−4.89]	0.064
10	−2.38	[−3.30]	−30.8	[−1.67]			0.014
10	−0.64	[−0.81]	72.9	[2.88]	−0.090	[−5.40]	0.062
15	−4.61	[−4.50]	−20.5	[−1.11]			0.004
15	−2.89	[−2.31]	81.6	[2.59]	−0.089	[−3.81]	0.036
20	−5.58	[−5.92]	−50.3	[−2.58]			0.019
20	−3.03	[−2.47]	101.5	[2.72]	−0.132	[−4.23]	0.071

Table 4: Post-Announcement Event

The cumulative post-announcement equity returns Δp_2^j (denominated in dollars and expressed in percentage points) for different event windows (WS) are regressed on a constant, the change in the risk premium $\left[\Sigma^G(w^n - w^o)\right]_j$ and the arbitrage risk $\left[\Sigma^G\Sigma^G(w^n - w^o)\right]_j$ of each stock j. Formally,

$$\Delta p_2^j = c + \alpha_2 \times [\Sigma^G(w^n - w^o)]_j + \beta_2 \times [\Sigma^G\Sigma^G(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for 2 years of weekly dollar stock returns for the period of July 1, 1998 to July 1, 2000. The event window size is chosen alternatively to extend over WS $= 3, 5, 7$ trading days starting on December 4, 2000. Panel A reports the coefficients for the entire sample, Panel B for only the added and deleted stocks and Panel C for the subsample of non-U.S. stocks. Robust and country clustered adjusted t-values are reported in parenthesis.

WS	c	[t]	α_2	[t]	β_2	[t]	R^2
Panel A: Post-Announcement Period (All Stocks, N=2291)							
3	0.67	[1.30]			0.023	[3.90]	0.037
3	0.12	[0.23]	−41.3	[−3.16]	0.053	[5.68]	0.085
5	1.90	[2.90]			0.047	[6.78]	0.138
5	1.13	[0.746]	−57.1	[−2.82]	0.089	[4.54]	0.155
7	1.40	[2.17]			0.033	[4.56]	0.069
7	0.61	[0.92]	−59.6	[−3.19]	0.077	[4.54]	0.087
Panel B: Post-Announcement Event (Only Added and Deleted Stocks, N=661)							
3	0.74	[1.79]			0.023	[7.08]	0.067
3	0.46	[1.01]	−24.3	[−0.148]	0.041	[3.17]	0.071
5	2.06	[3.51]			0.054	[6.26]	0.175
5	1.62	[2.22]	−37.8	[−1.39]	0.082	[3.03]	0.181
7	1.52	[2.22]			0.034	[6.81]	0.075
7	0.95	[1.33]	−48.5	[−2.23]	0.069	[3.88]	0.086
Panel C: Post-Announcement Event (Only Non-U.S. Stocks, N=1877)							
3	1.00	[2.14]			0.032	[4.06]	0.089
3	0.63	[1.33]	−25.1	[−2.08]	0.049	[4.66]	0.097
5	2.37	[4.65]			0.044	[3.76]	0.093
5	1.74	[3.00]	−42.1	[−2.10]	0.072	[4.11]	0.108
7	1.76	[2.84]			0.041	[3.30]	0.070
7	0.93	[1.24]	−56.2	[−2.37]	0.079	[3.75]	0.087

Table 5: Implementation Events

The cumulative event window equity returns Δp_3^j (denominated in dollars and expressed in percentage points) for different event windows (WS) are regressed on a constant, the risk premium change $[\Sigma^G(w^n - w^o)]_j$ and the arbitrage risk $[\Sigma^G\Sigma^G(w^n - w^o)]_j$ of each stock j. Formally,

$$\Delta p_3^j = c + \alpha_3 \times [\Sigma^G(w^n - w^o)]_j + \beta_3 \times [\Sigma^G\Sigma^G(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for 2 years of weekly dollar stock returns for the period of July 1, 1998 to July 1, 2000. The event windows start two days before the implementation day and extend over WS = 5, 7 trading days. Regression results for the first implementation event (November 30, 2001) are reported in Panel A for the entire sample, in Panel B for only the added and deleted stocks. Panels C and D provide corresponding results for the second implementation event (May 31, 2002). Robust and country clustered adjusted t-values are reported in parenthesis.

WS	c	[t]	α_3	[t]	β_3	[t]	R^2
Panel A: First Implementation Event (All Stocks, N=2291)							
5	2.47	[3.36]	21.4	[2.11]			0.026
5	1.42	[1.76]	−42.9	[−2.85]	0.055	[4.28]	0.061
7	2.93	[2.48]	12.2	[0.80]			0.005
7	1.70	[1.30]	−63.6	[−4.49]	0.064	[3.85]	0.036
Panel B: First Implementation Event (Only Added and Deleted Stocks, N=661)							
5	1.87	[2.07]	27.7	[2.56]			0.048
5	0.86	[0.98]	−40.7	[−1.76]	0.057	[3.25]	0.084
7	2.41	[1.71]	20.1	[1.38]			0.017
7	1.21	[0.88]	−61.1	[−2.53]	0.068	[3.97]	0.051
Panel C: Second Implementation Event (All Stocks, N=2291)							
5	−1.23	[−5.26]	−11.6	[−3.66]			0.015
5	−0.66	[−2.14]	23.0	[1.98]	−0.029	[−2.84]	0.035
7	−2.28	[−7.26]	−22.8	[−3.92]			0.032
7	−1.30	[−3.08]	37.4	[2.63]	−0.051	[−3.79]	0.067
Panel D: Second Implementation Event (Only Added and Deleted Stocks, N=661)							
5	−0.74	[−2.60]	−17.1	[−3.69]			0.044
5	−0.34	[−0.73]	9.9	[0.56]	−0.022	[−1.67]	0.057
7	−1.87	[−4.45]	−30.3	[3.57]			0.065
7	−1.19	[−2.13]	15.8	[0.75]	−0.038	[−2.60]	0.083

Table 6: Local versus Global Asset Pricing for Pre-Announcement Event

The cumulative pre-announcement equity returns Δp_1^j in stock j (denominated in dollars and expressed in percentage points) for different event windows (WS) are regressed on a constant, the change in the risk premium $[\Sigma^L(w^n - w^o)]_j$ of a local arbitrage portfolio, difference between the global and local risk premium change $[\Sigma^{Int}(w^n - w^o)]_j$, the arbitrage risks for the local arbitrage portfolio $[\Sigma^L \Sigma^L (w^n - w^o)]_j$ and the incremental international arbitrage risk to the global arbitrage risk $[\Sigma\Sigma^{Int}(w^n - w^o)]_j$. Formally,

$$\Delta p_1^j = c + \alpha_1^L \times [\Sigma^L(w^n - w^o)]_j + \alpha_1^{Int} \times [\Sigma^{Int}(w^n - w^o)]_j + \beta_1^L \times [\Sigma^L \Sigma^L (w^n - w^o)]_j + \beta_1^{Int} \times [\Sigma\Sigma^{Int}(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for 2 years of weekly dollar stock returns for the period of July 1, 1998 to July 1, 2000. The matrix Σ^L is obtained by setting to zero all stock covariances across countries to capture only within country arbitrage. Furthermore, $\Sigma^{Int} = \Sigma^G - \Sigma^L$ and $\Sigma\Sigma^{Int} = \Sigma^G\Sigma^G - \Sigma^L\Sigma^L$. The event window size is chosen alternatively to start WS $= 5, 10, 15, 20$ trading days prior to December 1, 2000. Panel A reports the coefficients for all stocks, Panel B only for added and deleted stock and Panel C only for non-U.S. stocks. Robust and country clustered adjusted t-values are reported in parenthesis. The last two columns report the significance level at which equality of the respective coefficients can be rejected.

WS	c	[t]	α_1^L	[t]	α_1^{Int}	[t]	β_1^L	[t]	β_1^{Int}	[t]	R^2	$\alpha_1^L = \alpha_1^{Int}$	$\beta_1^L = \beta_1^{Int}$
Panel A: Pre-Announcement Event (All Stocks, N=2291)													
5	1.49	[3.17]	−12.2	[−0.29]	52.6	[3.26]	−0.011	[−0.26]	−0.067	[−5.13]	0.101	0.207	0.249
10	−0.32	[−0.40]	22.8	[0.60]	90.8	[4.14]	−0.041	[−1.16]	−0.103	[−6.62]	0.091	0.058	0.079
15	−2.01	[−1.52]	120.0	[1.66]	99.3	[3.60]	−0.124	[−1.76]	−0.104	[−4.59]	0.052	0.761	0.760
20	−1.81	[−1.20]	144.8	[2.56]	130.5	[3.39]	−0.186	[−3.32]	−0.154	[−4.65]	0.121	0.797	0.545
Panel B: Pre-Announcement Event (Only Added and Deleted Stocks, N=661)													
5	1.14	[2.07]	11.1	[0.34]	48.9	[2.25]	−0.042	[−1.29]	−0.067	[−4.32]	0.151	0.412	0.555
10	−0.90	[−0.84]	25.0	[0.43]	71.9	[2.23]	−0.056	[−1.04]	−0.104	[−4.23]	0.131	0.315	0.293
15	−2.66	[−1.64]	219.2	[2.48]	80.1	[2.35]	−0.234	[−2.76]	−0.101	[−3.12]	0.084	0.068	0.071
20	−1.94	[−1.01]	170.8	[2.02]	102.6	[1.82]	−0.234	[−2.94]	−0.150	[−2.97]	0.183	0.334	0.217
Panel C: Pre-Announcement Event (Non-U.S. Stocks, N=1877)													
5	1.34	[2.70]	53.5	[0.85]	40.2	[2.35]	−0.147	[−0.53]	−0.058	[−4.26]	0.065	0.849	0.754
10	−0.87	[−0.98]	46.7	[0.75]	75.5	[2.67]	−0.171	[−0.47]	−0.091	[−4.77]	0.065	0.708	0.832
15	−2.88	[−1.99]	88.5	[0.83]	81.0	[2.27]	−0.107	[−0.21]	−0.089	[−3.25]	0.036	0.951	0.973
20	−3.11	[−2.12]	167.3	[1.92]	95.7	[2.31]	−0.361	[−0.89]	−0.127	[−3.65]	0.072	0.485	0.583

Table 7: Local versus Global Asset Pricing for Post-Announcement Event

The cumulative post-announcement equity returns Δp_2^j in stock j (denominated in dollars and expressed in percentage points) for different event windows (WS) are regressed on a constant, the arbitrage risks for the local arbitrage portfolio $[\Sigma^L\Sigma^L(w^n - w^o)]_j$ and the incremental arbitrage risk to the global arbitrage risk $[\Sigma\Sigma^{Int}(w^n - w^o)]_j$. Formally,

$$\Delta p_2^j = c + \beta_2^L \times [\Sigma^L\Sigma^L(w^n - w^o)]_j + \beta_2^{Int} \times [\Sigma\Sigma^{Int}(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for 2 years of weekly dollar stock returns for the period of July 1, 1998 to July 1, 2000. The matrix Σ^L is obtained by setting to zero all stock covariances across countries to capture only within country arbitrage. Furthermore, $\Sigma^{Int} = \Sigma^G - \Sigma^L$ and $\Sigma\Sigma^{Int} = \Sigma^G\Sigma^G - \Sigma^L\Sigma^L$. The event window size is chosen alternatively to extend over WS $= 3, 5, 7$ trading days starting on December 4, 2000. Panels A reports the coefficients for all stocks, Panel B only for added and deleted stock and Panel C only for non-U.S. stocks. Robust and country clustered adjusted t-values are reported in parenthesis. The last column reports the significance level at which equality of the respective coefficients can be rejected.

WS	c	[t]	β_2^L	[t]	β_2^{Int}	[t]	R^2	$\beta_2^L = \beta_2^{Int}$
Panel A: Post-Announcement Event (All Stocks, N=2291)								
3	0.71	[1.47]	0.019	[4.80]	0.032	[4.49]	0.076	0.005
5	1.93	[2.85]	0.049	[9.74]	0.044	[3.64]	0.139	0.674
7	1.44	[2.27]	0.029	[5.94]	0.042	[3.33]	0.073	0.208
Panel B: Post-Announcement Event (Only Added and Deleted Stocks, N=661)								
3	0.74	[1.76]	0.023	[8.84]	0.022	[2.78]	0.067	0.847
5	1.99	[3.35]	0.058	[14.85]	0.033	[2.18]	0.189	0.054
7	1.52	[2.20]	0.033	[8.16]	0.034	[2.67]	0.075	0.621
Panel C: Post-Announcement Event (Non-U.S. Stocks, N=1877)								
3	1.29	[4.41]	0.212	[1.42]	0.030	[4.07]	0.100	0.232
5	2.24	[5.30]	−0.033	[−0.15]	0.045	[3.86]	0.094	0.732
7	1.72	[3.23]	0.011	[0.264]	0.042	[3.21]	0.069	0.423

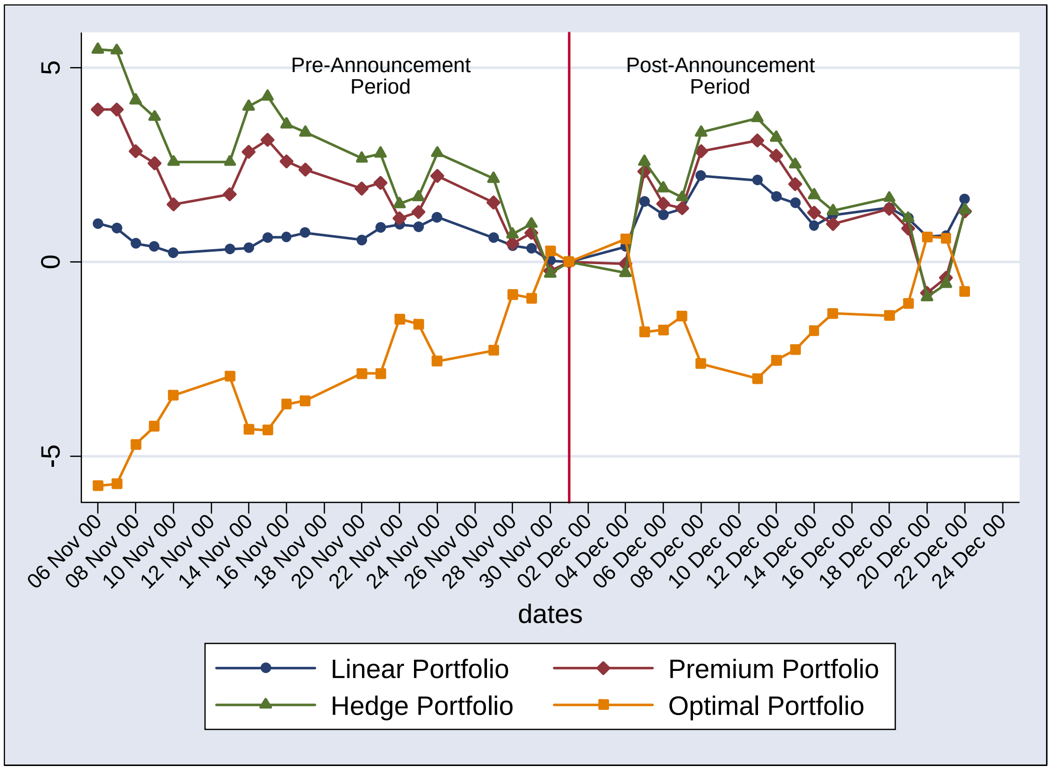

Figure 1: The cumulative returns around the announcement date of the index change are plotted for four self-financing portfolios with the sum of absolute weights normalized to unity. The 'linear portfolio' has weights proportional to the index weight changes $u = w^n - w^o$, the 'premium portfolio' has weights proportional to the risk premium changes Σu, the 'hedge portfolio' has weights proportional to the marginal arbitrage of risk $\Sigma\Sigma u$ of each stock, and the 'optimal portfolio' weights proportional to $\Sigma u - \theta\Sigma\Sigma u$ with $\theta = 0.001$.

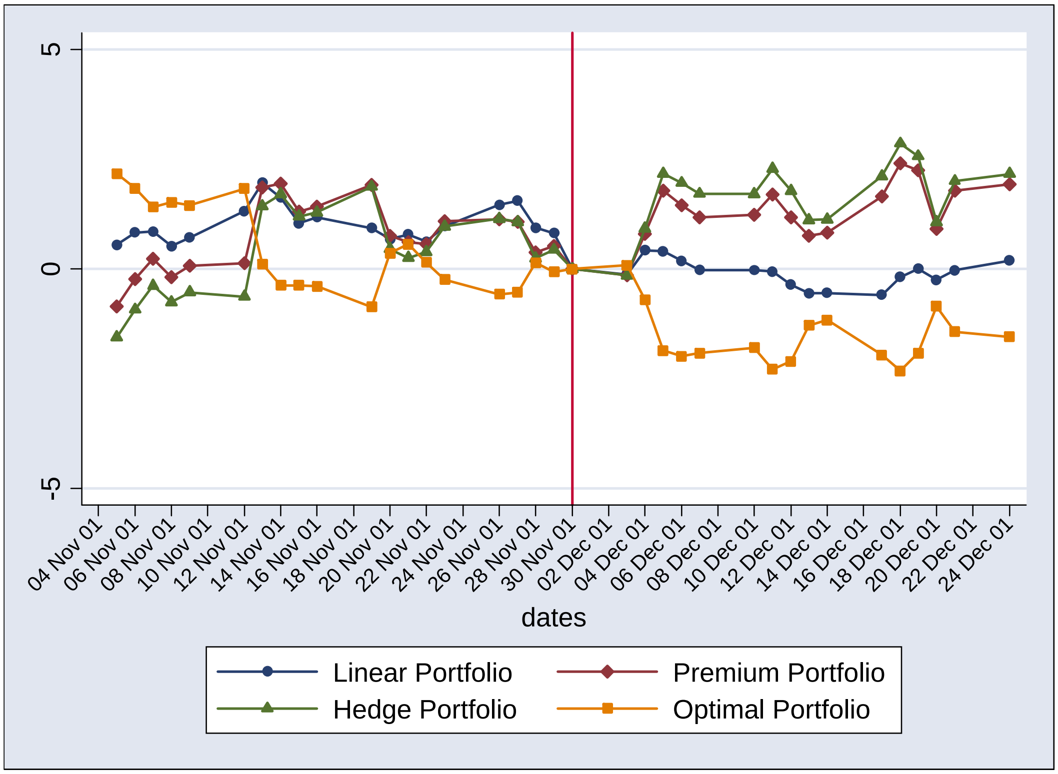

Figure 2: The cumulative returns around the first implementation date of the index change are plotted for four self-financing portfolios with the sum of absolute weights normalized to unity. The 'linear portfolio' has weights proportional to the index weight changes $u = w^n - w^o$, the 'premium portfolio' has weights proportional to the risk premium changes Σu, the 'hedge portfolio' has weights proportional to the marginal arbitrage of risk $\Sigma\Sigma u$ of each stock, and the 'optimal portfolio' weights proportional to $\Sigma u - \theta\Sigma\Sigma u$ with $\theta = 0.001$.

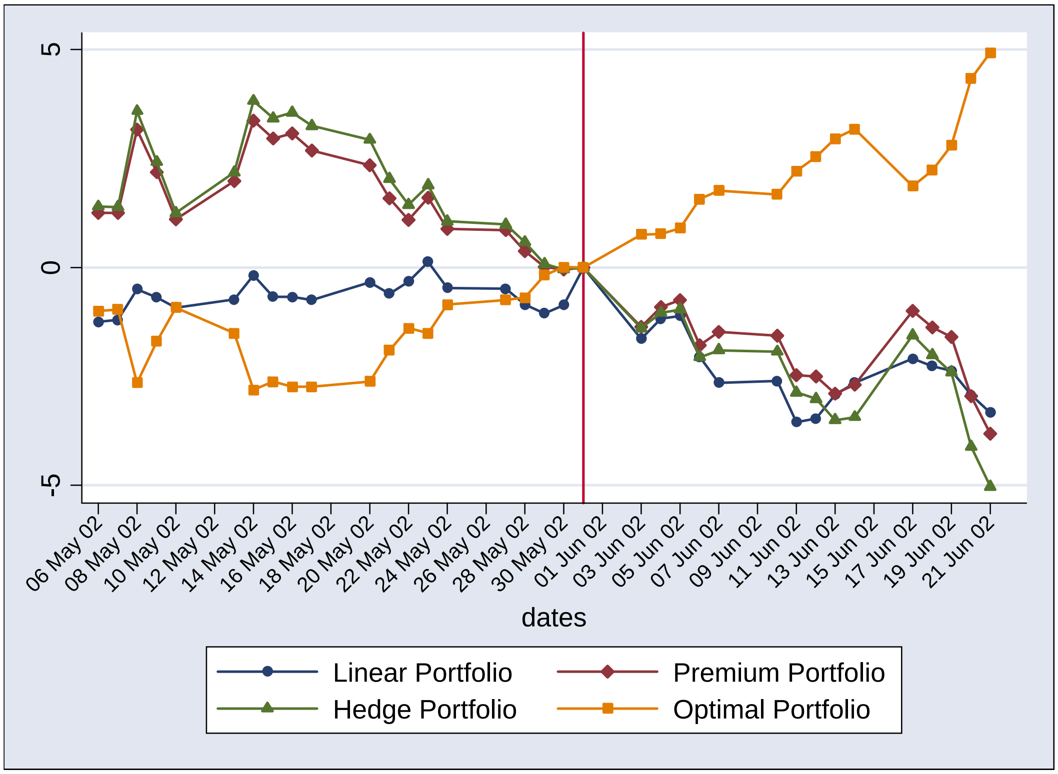

Figure 3: The cumulative returns around the second implementation date of the index change are plotted for four self-financing portfolios with the sum of absolute weights normalized to unity. The 'linear portfolio' has weights proportional to the index weight changes $u = w^n - w^o$, the 'premium portfolio' has weights proportional to the risk premium changes Σu, the 'hedge portfolio' has weights proportional to the marginal arbitrage of risk $\Sigma\Sigma u$ of each stock, and the 'optimal portfolio' weights proportional to $\Sigma u - \theta\Sigma\Sigma u$ with $\theta = 0.001$.

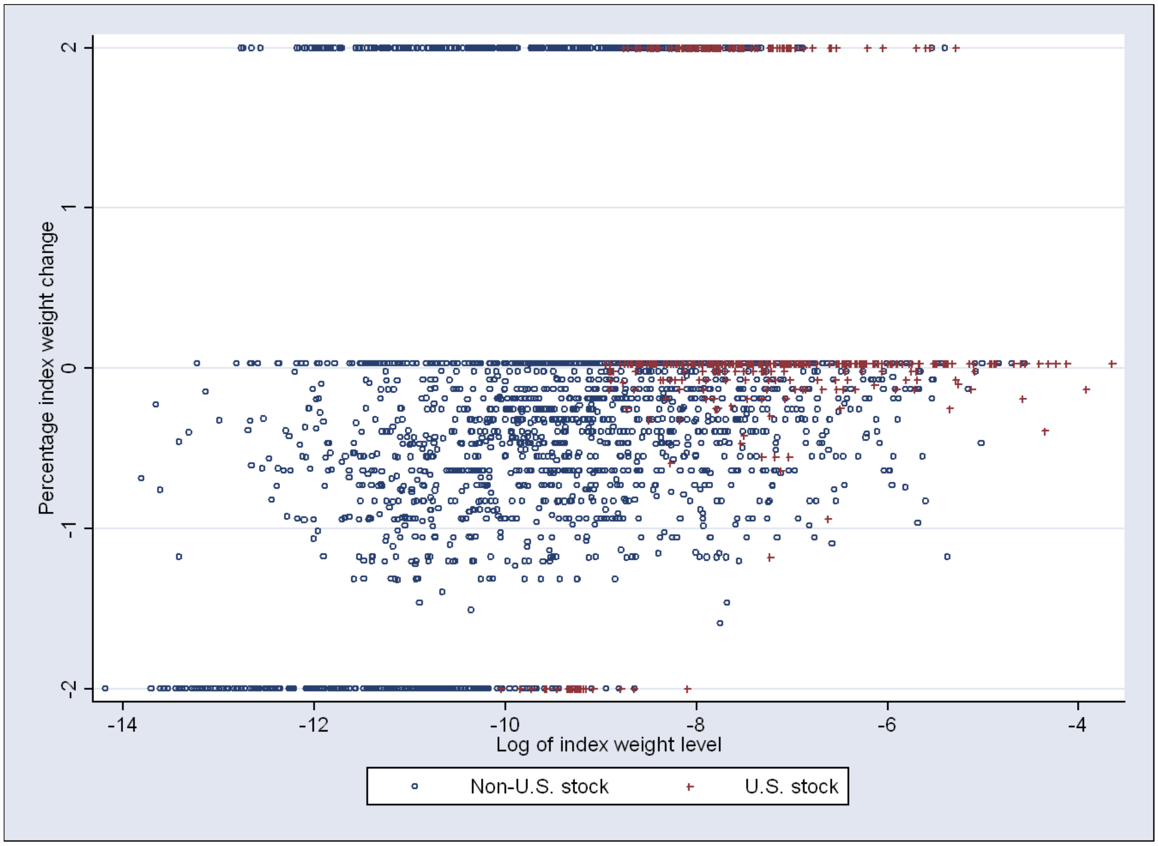

Figure 4: The percentage weight change for U.S. and non-U.S. stocks is plotted as a function of the log of the level of the old weight in the index (or the new weight in the case of stock additions).

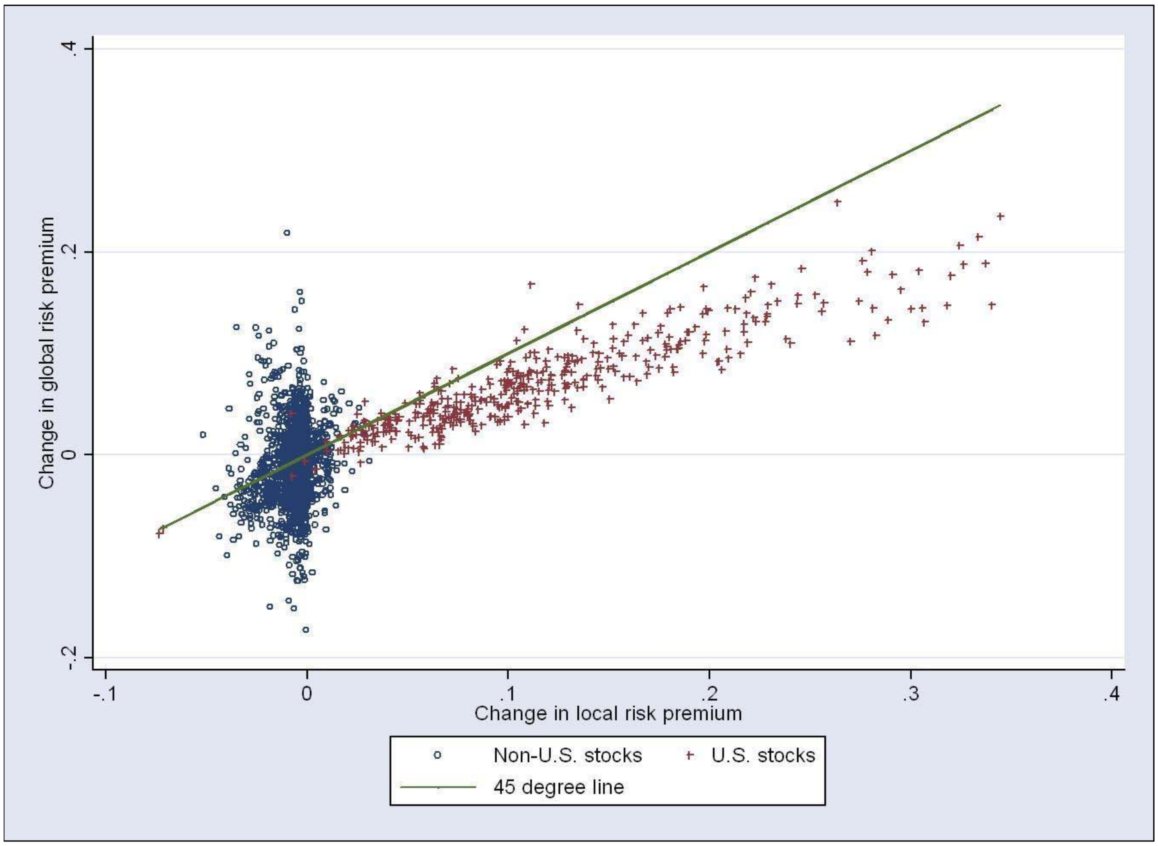

Figure 5: Risk premium change $[\Sigma^L(w^n - w^o)]_j$ of stock j under local asset pricing (market segmentation) are plotted against the risk premium change $[\Sigma^G(w^n - w^o)]_j$ of the same stock under global asset pricing (market integration).

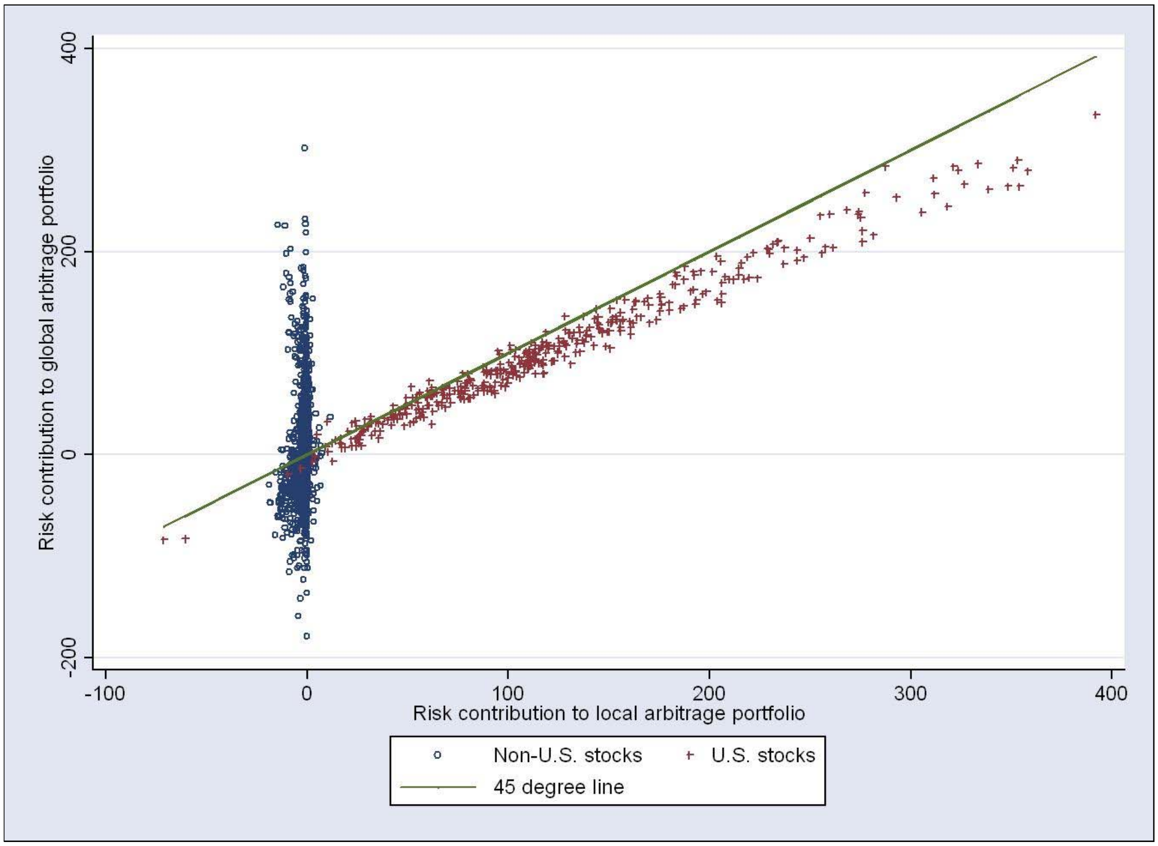

Figure 6: The arbitrage risk contributions $[\Sigma^L \Sigma^L (w^n - w^o)]_j$ of individual stocks j to local arbitrage portfolios composed only of local stock (x-axis) are plotted against the arbitrage risk contributions $[\Sigma^G \Sigma^G (w^n - w^o)]_j$ of the same stock to a global arbitrage portfolio of all stocks (y-axis).

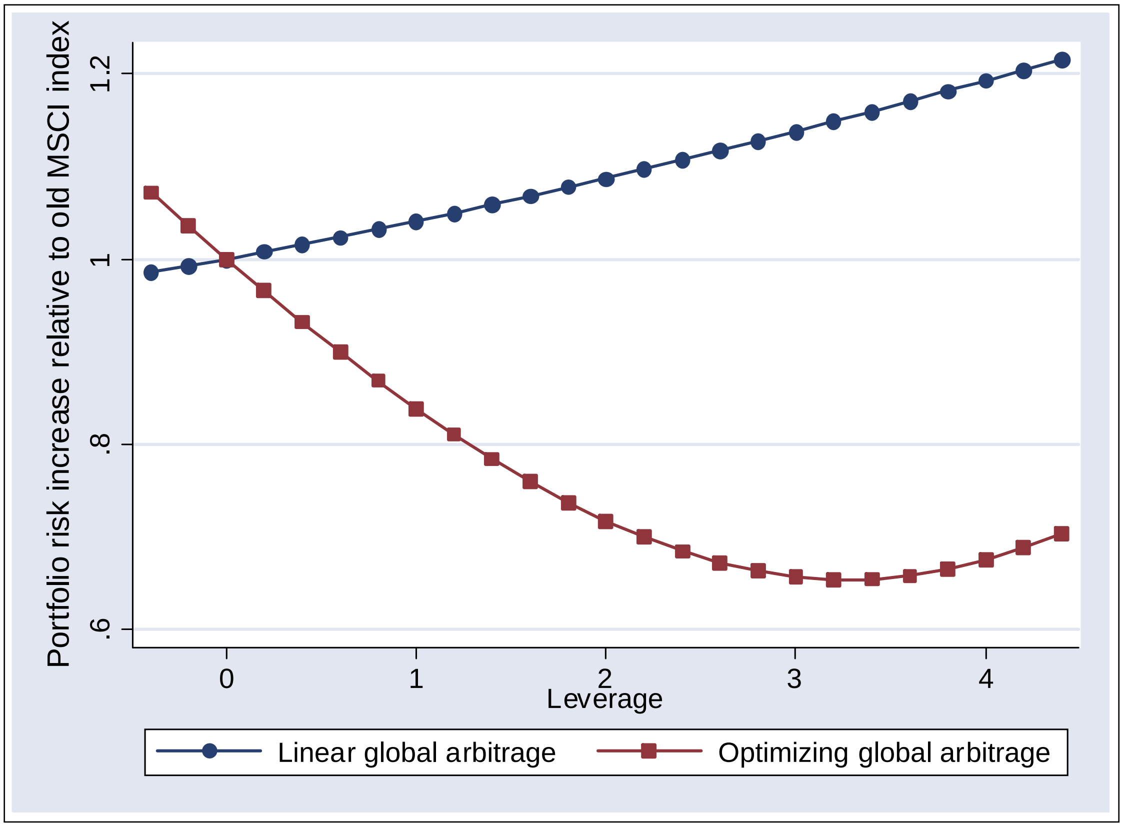

Figure 7: The total risk of an arbitrage portfolio relative to the total portfolio risk of the old MSCI index is plotted for two alternative arbitrage strategies: *Linear global arbitrage* represents a shift (with leverage factor κ) into the linear portfolio ϖ^{Lin} and *Optimizing global arbitrage* represents a shift (under the same leverage) into the optimal portfolio ϖ^{Opt} with parameter $\theta = 0.001$.