

Global versus Local Asset Pricing: A New Test of Market Integration

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Should capital cost calculations be based on a global or local market benchmark? The December 2000 redefinition of the Morgan Stanley Capital International (MSCI) global equity index was a natural experiment addressing this question. It is argued that this event triggered a portfolio shift (by index funds) large enough to affect the residual asset supplies constituting the global and local market benchmarks of all actively managed capital. Changes in the market benchmarks imply distinct and predictable changes to global and local stock betas. Exploring whether global or local beta changes best explain the cross-section of event returns reveals that stocks in developed markets are priced globally and not locally. (*JEL* G11, G14, G15)

Should capital cost estimation be based on a global or local market portfolio? If asset markets are integrated, then a world market portfolio represents the appropriate benchmark, whereas market segmentation implies risk pricing to a national market benchmark. There seems to be no consensus on the answer to this question despite its practical ramifications for corporate finance. This article provides a new perspective on the issue based on evidence from a natural asset pricing experiment. The revision of the global Morgan Stanley Capital International (MSCI) equity index in December 2000 dramatically changed the country representation in the index with a median absolute percentage weight change of 42% at the country level. It is argued here that this substantial index change for a large amount of index-tracking capital modified global and local market benchmarks that are based on actively managed assets. Such benchmark changes imply distinct and predictable changes to all global and local equity betas. A careful analysis of the event returns around the announcement of the index change can therefore reveal the degree of global versus local asset pricing. Global beta changes should, *ceteris paribus*, account for the event returns under a global market benchmark, whereas a local

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benchmark is appropriate if local beta changes best explain the cross-section of returns. The main contribution of this article is to show that global beta changes best account for the event returns of the MSCI index change.

Finance research has long acknowledged the importance of a “correct market identification” for testing asset pricing models. Roll (1977) famously expressed skepticism about the possibility of identifying the market benchmark. Testing asset pricing models involves a joint hypothesis about the market benchmark and the model itself, and the joint hypothesis problem tends to make any model rejection difficult to interpret. This article finds a way around the problem of an unobservable market benchmark by adopting a “difference in differences” approach: Instead of identifying the market benchmark, this article focuses on a large *exogenous change* to the market benchmark. Such a benchmark change may be easier to identify than the correct market benchmark itself.

A market benchmark change occurs if a group of investors does not optimize its portfolio choice and engages in a large (exogenous) portfolio reallocation. Index investors constitute one such group. Their capital is invested according to fixed portfolio weights and—unlike the market benchmark capital—is not subject to continuous portfolio optimization.¹ Equity indices periodically undergo small revisions, but in rare cases the index redefinition is substantial and generates a massive portfolio reallocation. Such a portfolio shift by index investors represents a shock to the market benchmark because it changes the amount of residual equity capital held by the optimizing investors.² Moreover, a market benchmark change induced by an index revision is straightforward to measure. It is proportional to the index weight change $w^n - w^o$, where w^n and w^o denote the vector of new and old weights, respectively.

The asset pricing implications of a market benchmark change are fairly intuitive. The capital asset pricing model (CAPM) beta of a stock is proportional to the covariance of its return with the market benchmark return. Hence, a change in the market benchmark implies a corresponding change in stock betas. A stock’s beta change modifies the discount rate and translates into a stock price change. A major index revision can therefore serve as a natural experiment to test an asset pricing model “in differences.” Such a test is robust to the existence of other unspecified risk factors as long as these are uncorrelated with the index change. Indeed, such additional risk factors should not matter for price effects over very short event intervals around the index change.³

¹ All shares held for control reasons, such as government stakes or family holdings, are also not part of the market benchmark definition. A further reduction of the “investable” part of the market might result from investor concerns about governance (Leuz, Lins, and Warnock 2009).

² For simplicity, I assume here that the index investors and the optimizing portfolio investors are distinct groups and that investors fall into either of the two groups.

³ MSCI weight changes were determined by a stock’s “free float” relative to its market capitalization—a ratio influenced mostly by strategic, long-term family and/or state ownership in a stock. The latter variables are not

My empirical strategy is based on the redefinition of the MSCI All Country World Index (ACWI) announced in December 2000. The weight changes followed from the adoption of so-called free-float weights, and they led to substantial and simultaneous index weight changes in 2566 stocks worldwide. The magnitude of the weight changes and the large amount of index capital that tracks the MSCI global equity index make this event an exceptional market benchmark change. The global nature of the MSCI index change enables examination of various hypotheses about market segmentation. The index revision influences global beta changes differently than it does local market betas or betas corresponding to particular market segments. Under the hypothesis of market integration, global beta changes should fully explain event returns. On the other hand, market segmentation allows for the explanatory power of beta changes that correspond to market segments.

Testing for market integration therefore requires specific hypotheses about the nature of potential market segmentation. I focus on four hypotheses characterizing the most plausible dimensions of segmentation. First, equity markets might be segmented along national markets (defined by a stock's primary listing). If so, then national market beta changes should have exclusive explanatory power for the price adjustment. The data provide little support for this hypothesis, and they reject the hypothesis of market segmentation along national markets. In contrast, the hypothesis of a single global market benchmark change cannot be rejected. Second, market segmentation may exist between developed and emerging markets. Yet, such a market segmentation does not explain event returns either: Index weight changes of developed market stocks show a significant return impact on emerging market stocks that is in line with their global beta changes. Third, I examine whether global market integration (in terms of risk pricing) is more pronounced for cross-listed emerging market stocks than for those without a cross-listing. Fourth, emerging market stocks are partitioned into the 50% most liquid and least liquid stocks to test for market segmentation along the liquidity dimension. Cross-listing (in the United States or in the United Kingdom) and more liquidity both are found to be associated with global risk pricing. Emerging market stocks without a cross-listing and with low liquidity show evidence of market segmentation.

This article contributes to the literature on international asset pricing from a new methodological angle. [Karolyi and Stulz \(2003\)](#) survey the literature on global versus local asset pricing. Empirical work has increasingly tested the world CAPM in a conditional setting with time-varying expected returns, variances, and covariances—as studied by [Harvey \(1991\)](#), [Chan, Karolyi, and Stulz \(1992\)](#), [De Santis and Gerard \(1997\)](#), and [Zhang \(2006\)](#). Conditional

endogenous at the business cycle frequency for which risk factors are constructed. Also, empirical asset pricing has not yet documented significant linkages between risk factors and equity ownership structure. The latter is different from shareholder rights, which [Gompers, Jshii, and Metrick \(2003\)](#) find related to excess performance in the 1990s in an international sample of 1,500 large stocks.

models typically feature many free parameters, and the corresponding loss of statistical power may explain the weak empirical support for the world CAPM model. Dynamic models, which nest both polar cases of market integration and segmentation, may also depend on unobservable state variables. For example, [Carrieri, Errunza, and Hogan \(2007\)](#) estimate an international asset pricing model in which the correlation of the domestic market return with a set of (most correlated) “eligible” foreign assets serves as a proxy for market integration. Their framework requires a choice of particular global assets that best replicate country returns. In contrast, my identification strategy does not depend on unobservable state variables, and neither do I require any identification assumption (in levels) about the global market benchmark or its segmented components. This may explain why I find strong evidence that the global CAPM beta is, indeed, a priced factor.

The increasing role of index investment has created much interest in its asset pricing implications. [Brennan and Li \(2008\)](#) argue that index tracking by institutional investors creates an investment bias toward S&P500 stocks, which reduces the residual supply of S&P500 index risk. Accordingly, the authors find evidence that stocks whose returns covary more with the idiosyncratic component of the S&P500 return have significantly lower returns in recent periods.⁴ This suggests that index investment is significant enough to modify equilibrium returns.

Many studies document the stock price impact of index inclusions and exclusions. These event studies initially focused on individual stocks, showing that index inclusions increase share prices and that exclusions decrease them.⁵ However, individual index inclusions or exclusions do not substantially modify the market benchmark, so the focus has been on temporary “price pressure” and its long-run reversal. Similarly, a broader literature on “liquidity effects” assesses whether demand shocks correlate with individual stock price returns.⁶ A marketwide index change has different implications. More specifically, robustness tests suggest that price pressure or liquidity effects do not account for the price dynamics of the MSCI index revision.

Finally, I highlight a limitation of the analysis and an additional contribution. The event study approach gives only a snapshot of the degree of financial integration around the event date. Therefore, no issues relating to the time

⁴ For related evidence on low returns for stocks with high idiosyncratic volatility relative to the Fama and French (1993) model, see [Ang et al. \(2006\)](#).

⁵ See [Shleifer \(1986\)](#), [Kaul, Mehrotra, and Morck \(2000\)](#), and [Hegde and McDermott \(2003\)](#).

⁶ 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 \(2001\)](#) show 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 \(1999\)](#) document 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.

variation in financial integration can be addressed and this is clearly a limitation.⁷ But this event study also contributes to our understanding of speculative trading: The index revision could be anticipated by speculators, and theoretical considerations suggest that their hedging demand might have an additional event return effect. Therefore, any estimation of the fundamental return effect of beta changes should control for the transitory price impact of hedging demands; hence, I include a specific hedging term in all test specifications. Accounting in this way for speculative hedging yields new insights into the structure of international risk arbitrage. I explicitly derive the optimal arbitrage strategy of risk-averse speculators and also document the exact price impact of their trading.

The article proceeds as follows. Section 1 provides an intuitive discussion of the empirical strategy. A formal dynamic model of multi-asset arbitrage is presented in Section 2, together with its testable predictions. 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 4 provides the evidence; the hypotheses of global versus local asset pricing are tested in Section 4.3, and segmentation between emerging and developed markets is examined in Sections 4.4 and 4.5. Section 5 discusses robustness issues, and Section 6 concludes.

1. A Heuristic Discussion

1.1 New risk premiums due to a large-scale index change

The market benchmark is defined as the total tradable capital held by optimizing investors. This excludes index investors who hold equity stakes passively according to a vector of index weights w^o . Let parameter C scale these weights to the total (dollar amount of) index capital. The market benchmark adjusted for the passive index investment is then $S^o = \bar{S} - Cw^o$, where \bar{S} denotes the (hypothetical) market benchmark if there were no index investors. For simplicity, I normalize the benchmark asset supply to 1. Thus, benchmark weights are standardized as $\tilde{S}^o = S^o/1S^o$. For a return vector $r = (r_1, r_2, \dots, r_n)$, the market benchmark return r_m^o can be written as

$$r_m^o = r\tilde{S}^o. \quad (1)$$

The beta of stock j follows as

$$beta_j^o = \frac{Cov(r_j, r_m^o)}{Var(r_m^o)} = \frac{1}{Var(r_m^o)} [\Sigma\tilde{S}^o]_j, \quad (2)$$

where Σ denotes the $n \times n$ covariance matrix of asset returns r and the subscript j represents the j th element of the $n \times 1$ row vector.

⁷ Bekaert et al. (2009) document a decrease of equity market segmentation over the 1980s and 1990s, for both developed and emerging countries, and explore the determinants of this evolution. See also Bekaert and Harvey (1995) and Carrieri, Errunza, and Hogan (2007) for evidence on the time variation of market integration.

Next, consider the asset pricing implications of an index change from weights w^o to w^n . First, I assume that the benchmark volatility $Var(r_m)$ does not change. Let $\vartheta = C/1S^o$ represent the ratio of index tracking capital to total benchmark capital. The beta change is then simply

$$\Delta beta_j = beta_j^n - beta_j^o = \frac{Cov(r_j, r_m^n - r_m^o)}{Var(r_m^o)} = -\frac{\vartheta}{Var(r_m)} [\Sigma(w^n - w^o)]_j. \quad (3)$$

The index change leaves the cash flow expectations of any stock unchanged, so that only the discount factor changes. The corresponding stock return effect Δr_j around the announcement of the index change is proportional to $[\Sigma(w^n - w^o)]_j$; formally,

$$\Delta r_j \sim [\Sigma(w^n - w^o)]_j. \quad (4)$$

In the general case of a change in the benchmark volatility $Var(r_m)$, the equilibrium equity risk premium also changes. It is straightforward to show that a mean–variance framework implies a proportional change in the equity market premium and so the ratio $[\mathcal{E}(r_m) - r_f]/Var(r_m)$ remains constant. Hence, the cross-sectional asset return effect is proportional to $[\Sigma(w^n - w^o)]_j$ even if the benchmark volatility changes under the index revision.⁸

1.2 Controlling for the speculative arbitrage dynamics

Speculative trading prior to the public announcement of the index change may accelerate the price adjustment to the new discount rates. But such speculative trading may also generate additional confounding price effects if speculators are risk averse and trade against less informed liquidity suppliers. This article develops a new model for the resulting speculative price dynamics.⁹ Consider the optimal arbitrage strategy of a risk-averse mean–variance investor who is privately informed about the index revision. His optimal arbitrage position is shown to feature two distinct terms. First, it has a *return-seeking component* proportional to the vector of expected excess returns given by $\Sigma(w^n - w^o)$. This component is due to the change in the market benchmark from S^o to S^n and the corresponding change in stock betas. Second, it features a *risk-hedging component* proportional to the stock-specific marginal arbitrage risk contribution; the latter is shown to be proportional to $\Sigma \Sigma(w^n - w^o)$. Optimization in the mean–variance space requires arbitrageurs to choose a portfolio that optimizes the trade-off between expected arbitrage returns and marginal arbitrage risk in each stock. The optimal arbitrage strategy reduces

⁸ A change in the discount factor may alter the price of growth stocks differently from value stocks because of a different intertemporal cash flow pattern. However, such differences are likely to be of second order and are ignored in the subsequent analysis.

⁹ The model is related to Greenwood (2005) and nests his model as a special case if there are no uninformed liquidity suppliers. See also Greenwood and Vayanos (2008) for a setting with uninformed liquidity suppliers.

(increases) the portfolio weight of stocks with a positive (negative) marginal arbitrage risk contribution. During implementation of the arbitrage strategy, both the return-seeking and the risk-hedging components may simultaneously influence stock prices and generate a cross-sectional return effect given by

$$\Delta r_j \sim [\Sigma(w^n - w^o)]_j - k [\Sigma \Sigma(w^n - w^o)]_j. \quad (5)$$

for some $k > 0$.

An additional contribution of this article is to show that the new model correctly characterizes the cross-section of asset returns around the MSCI announcement event. The return-seeking and risk-hedging components both show the correct regression sign, and they both have a (statistically and economically) highly significant price effect. The importance of the risk-hedging component directly reflects the speculators' limited risk tolerance.

1.3 Testing for market integration

In international finance, the issue of market integration is typically addressed as global versus local (i.e., national) asset pricing. In the MSCI event setting, the portfolio shift by index investors modifies both the global and the local market benchmark. Under market integration, domestic and foreign stocks should, *ceteris paribus*, contribute equally to the global market benchmark and also to a benchmark change. Alternatively, market segmentation by country implies that only domestic stocks constitute the market benchmark and that it is only the change in the local benchmark that generates a return effect.

The test strategy can best be illustrated for market segmentation along national stock markets. Here, the market integration test consists of isolating local from global beta changes. The global covariance matrix Σ^G of all index stocks can be decomposed into a matrix Σ^L featuring nonzero covariance elements only between stocks in the same country and a complementary matrix $\Sigma^{Int} = \Sigma^G - \Sigma^L$ featuring only cross-country (international) stock covariances and zeros otherwise. The global beta changes are proportional to $\Sigma^G(w^n - w^o)$, and local beta changes are proportional to $\Sigma^L(w^n - w^o)$. Global asset markets are segmented into local markets if the difference between the global and local beta changes given by $\Sigma^{Int}(w^n - w^o)$ does not help explain event returns. Alternatively, market integration 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.

The methodology is general enough to test for a range of alternative market segmentation hypotheses. For example, equity markets might be integrated within the group of developed market stocks, while segmentation prevails between developed and emerging market stocks. For this hypothesis,

the appropriate decomposition of the global covariance matrix consists of a matrix Σ^H containing only covariance elements, where both stocks are developed market stocks or both are emerging market stocks, and a complementary (cross-hemisphere) matrix Σ^{CH} consisting of covariance elements between developed and emerging market stocks. Segmentation then implies that the term $\Sigma^{CH}(w^n - w^o)$ does not contribute to the event returns. An even finer partition of the matrix Σ^{CH} —for example, by a cross-listing criterion for the emerging market stock or by its liquidity—yields a test about the global market integration of specific groups of emerging market stocks.

A few similarities and differences to the existing literature can be highlighted. Previous empirical work on the degree of international equity market integration used capital market liberalization as the identifying event to measure risk premium changes (Bekaert and Harvey 2000; 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 much shorter event window. Moreover, the exogeneity assumption about the index change in this article is easier to defend than that of a liberalization policy, which may simultaneously affect future company cash flows. Other work has focused on cross-listing events in U.S. markets as a trigger for risk premium changes (Foerster and Karolyi 1999). Yet, similarly to equity issues, cross-listing decisions may be related to asymmetric information about cash flow prospects and therefore might not qualify as purely exogenous events.

2. Theory and Hypotheses

2.1 Model assumptions

This section develops a simple limits-to-arbitrage model that allows me to analyze the return effects of demand shocks in a multi-asset setting. A set of n financial assets are traded in regular intervals Δt . The market characteristics are summarized as follows.

Assumption 1 (Market structure, asset supply, and liquidation value). The financial market allows simultaneous trading in risk assets $i = 1, 2, 3, \dots, n$. Trading takes place over the time interval $[0, T)$ at equally spaced time points $t = 0, \Delta t, 2\Delta t, 3\Delta t, \dots, T - \Delta t$, with $\Delta t = T/N$. Liquidation occurs at time T at a price

$$p_T = \mathbf{1} + \sum_{t=\Delta t}^T \varepsilon_t, \quad (6)$$

where ε_t denotes serially uncorrelated mean-zero innovation learned by all market participants at time t . The innovations ε_t have a constant covariance matrix $\mathcal{E}_{t-\Delta t}(\varepsilon_t \varepsilon_t') = \Sigma \Delta t$. The market benchmark representing the actively traded supply of risky assets is given by a vector \tilde{S}^o . At time t_u , a demand

shock $u = \vartheta(w^n - w^o)$ occurs owing to the rebalancing of index investors from old index weights w^o to new index weights w^n , where ϑ denotes the ratio of index capital to total benchmark capital. The initial ($t = 0$) expected liquidation value of all assets is normalized to the unit vector $\mathbf{1}$.

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

A new model feature (compared to the study in Greenwood 2005) is the introduction of liquidity-supplying agents. These are the potential counterparties to the arbitrageurs who seek a net arbitrage position. The arbitrage opportunity is further embedded in the assumption that liquidity suppliers learn about the exogenous liquidity shock only with a delay. It is then shown that the existence of less informed liquidity suppliers generates additional hedging effects for the cross-sectional price patterns of event returns. Assumption 2 characterizes the investment behavior of these two types of market participants.

Assumption 2 (Risk arbitrageurs and linear liquidity supply). A unit interval of market participants can be grouped into a set $[0, \lambda]$ of risk arbitrageurs (or speculators) and a set of liquidity suppliers $(\lambda, 1]$. Arbitrageurs have a constant absolute risk aversion (CARA) utility, a risk aversion parameter ρ , and access to a riskless asset of zero return. Their optimal demand vector is then

$$x_t^A = (\rho \Sigma \Delta t)^{-1} \mathcal{E}_t^A(p_{t+\Delta t} - p_t), \quad (7)$$

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

$$x_t^L = \gamma \mathcal{E}_t^L(p_{t+\Delta t} - p_t), \quad (8)$$

where \mathcal{E}_t^L characterizes the expectations of the liquidity suppliers.

The arbitrageurs are optimizing agents who maximize CARA utility over their short investment horizon Δt . The Greenwood (2005) framework is nested and recovered for a parameter $\lambda = 1$ when only arbitrageurs constitute the market.¹⁰

¹⁰ Formally, Greenwood (2005) builds on the asset pricing framework in Hong and Stein (1999) 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.

The liquidity suppliers represent a new (ad hoc) addition to the model. “Representative agent” models appear generally inconsistent with regard to existing evidence of steep demand curves for individual stocks (Petajisto 2009). Limited market participation and heterogeneous cash flow expectations by liquidity suppliers tend to generate lower supply elasticities (lower γ) that are more plausible and thereby justify the reduced-form assumption in Equation (4). The expectation term $\mathcal{E}_t^L(p_{t+\Delta t})$ may more generally be interpreted as the (shifting) parameter of an asset supply function for which the price $p_t = \mathcal{E}_t^L(p_{t+\Delta t})$ yields a zero net (aggregate) asset supply. If differences in cash flow expectations determine the liquidity supply in each stock, then the latter should show little or no dependence on the covariance structure of returns—unlike the asset demand of the arbitrageurs. Moreover, any liquidity supply based on heterogeneous cash flow expectations generates a constant supply elasticity γ even as the trading intervals Δt become shorter.

An alternative liquidity formulation could assume that the asset supply does not occur stock-by-stock, but depends on the joint return covariance risk Σ according to

$$x_t^L = (\rho^L \Sigma)^{-1} \mathcal{E}_t^L(p_{t+\Delta t} - p_t), \quad (9)$$

where ρ^L can be interpreted as a risk aversion parameter of the liquidity suppliers. A liquidity supply of this nature produces a very different price dynamics because speculators do not hedge their arbitrage positions in equilibrium. Intuitively, if liquidity suppliers and speculators share the same beliefs about portfolio risk, then an exchange of hedging positions cannot provide mutual trading benefits. The analysis here focuses on the more interesting case for which hedging occurs.¹¹

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 diagonal matrix Γ , where stock-specific liquidity supply elasticities feature as the diagonal elements. None of the model insights depend on this modification.¹² 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. In order to keep the model simple, I do not develop a rational expectations equilibrium in which the liquidity suppliers rationally anticipate the possibility of an index revision and

¹¹ See the Online Appendix at www.haraldhau.com for a solution to the alternative model setup.

¹² It can be shown that stock-specific liquidity differences do not alter the return effect of the premium change, which is still proportional to Σu . Intuitively, arbitrageurs modify 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 because lower speculative positions for low-liquidity stocks require also smaller hedge positions. The arbitrage risk effect on returns is proportional to $\Gamma \Sigma \Sigma u$.

infer its likelihood from the trading behavior of arbitrageurs. Instead, beliefs about the supply shocks are set exogenously. The arbitrageurs learn about the index weight change at time t_A , whereas liquidity suppliers learn about it only with a delay at time $t_L > t_A$. Such an exogenous formulation of beliefs may be justified by recognizing that the index revision represents a rare event and is thereby not subject to dynamic learning through prices. Moreover, extending the model to a setting of dynamic learning by the liquidity providers requires new assumptions about the extent to which “noise trading” prevents complete inference. Intuitively, less noise facilitates dynamic inference and tends to limit the liquidity supply. But limited liquidity supply is already a model feature that is captured flexibly by the parameter λ . The limit case *without* noise trading corresponds to the fully revealing equilibrium without liquidity supply ($\lambda = 1$). The model setup therefore appears sufficiently general for an empirical characterization of the market event. Rational inference by the liquidity suppliers does not add any relevant aspects that are not already nested in the model parameterization.

Assumption 3 (Information structure and beliefs). Initially, both the arbitrageurs and the liquidity suppliers believe that the asset supply remains constant at S . At time t_A , arbitrageurs learn about the net supply changes from \tilde{S} to $\tilde{S} - u$. Arbitrageurs correctly anticipate that liquidity suppliers learn about the net supply soon after, at time $t_L > t_A$. The net supply changes occur at time $t_u > t_L$, and all assets are liquidated at time $T > t_u$ with a price p_T .

All arbitrageurs are assumed to learn about the demand shock u at the same time t_A and immediately seek a speculative position. In this context, it is appropriate to discuss the market mechanism for market clearing. In a modern open limit order book, orders are executed sequentially against the increasing price schedule of the liquidity suppliers. Order volume executed first earns the largest informational rents. However, if a batch auction is assumed as the market mechanism at t_A , then all order execution occurs at the uniform price p_{t_A} and thus the speculators’ informational rents are competed away. Empirically, anticipation of the supply shock and implementation of speculative positions are likely to stretch over many days and the empirical part of the analysis will account for this.

2.2 Speculative price dynamics

The trading process can be divided into four distinct phases. For $0 \leq t < t_A$, risk arbitrageurs and liquidity suppliers both assume a constant net asset supply of \tilde{S}^o . In the interval $t_A \leq t < t_L$, only the arbitrageurs know about the future change of the net supply. In this phase, the asset valuations of arbitrageurs and liquidity suppliers diverge. This valuation difference disappears at time t_L , when liquidity suppliers share the arbitrageurs’ information about the new future risk premiums. The last phase, $t_u \leq t < T$, is marked by the new risk

premiums, change due to the change in actively traded asset supply until asset liquidation occurs at time T . The equilibrium price process p_t is solved using backward induction: The solution starts from the market-clearing condition at the last trading period $T - \Delta t$ and then proceeds through repeated substitution to progressively earlier stages of the price process. The derivation is provided in the Appendix.

Of particular interest is the excess return at time t_A , when arbitrageurs learn about the demand shock. The following proposition characterizes this excess return.¹³

Proposition 1 (Excess returns of the speculative position buildup). Arbitrageurs acquire speculative positions at time t_A , when they learn about the future asset demand shock $u = \vartheta(w^n - w^o)$ occurring at time t_u due to a revision of old index weights w^o to new index weights w^n . The excess return at time t_A is positively proportional to the premium change $\Sigma(w^n - w^o)$ and negatively proportional to the arbitrage risk term $\Sigma \Sigma(w^n - w^o)$, where Σ denotes the covariance matrix of asset returns. Formally, the following linear approximation is obtained:

$$\Delta r_{t_A} \approx \alpha_1 \Sigma(w^n - w^o) + \beta_1 \Sigma \Sigma(w^n - w^o), \quad (10)$$

with $\alpha_1 = \frac{\rho}{\lambda} \vartheta(T - t_u) > 0$ and $\beta_1 = -(1 - \lambda) \gamma \left(\frac{\rho}{\lambda}\right)^2 \vartheta(T - t_u)(t_L - t_A) < 0$.

Proof. See the Appendix.

In the baseline case of the Greenwood model with $\lambda = 1$, the announcement price effect simplifies to the single term $\alpha_1 \Sigma(w^n - w^o)$ because $\beta_1 = 0$. This term represents the fundamental valuation effect of the beta changes of all stocks due to the asset demand shock $w^n - w^o$. As noted in Section 1.1, the term $\Sigma(w^n - w^o)$ can be related to the changes of the stock betas:

$$\Delta \text{beta} = -\frac{\vartheta}{\text{Var}(r_m)} \Sigma(w^n - w^o), \quad (11)$$

where $\text{Var}(r_m)$ denotes the market volatility (assumed here to be constant) and $\vartheta = C/1S^o$ the ratio of MSCI index-tracking capital to total benchmark capital. As the volume of index-tracking capital becomes large relative to the benchmark capital (of actively traded equity), any index weight change can have a non-negligible effect on stock betas. The coefficient α_1 includes the factor $T - t_u$, which represents the time that elapses between the net supply shock and the terminal cash payout. Over this duration, the stock betas are changed. The asset price effect captured by $\Sigma(w^n - w^o)$ is referred to as

¹³ The normalization of the liquidation price vector to $\mathbf{1}$ implies that any price change translates into an (approximately) equally large event return.

the *risk premium change* because it is proportional to the beta change and the change in the discount factor. 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. I emphasize that the price effect does not depend on a correct specification of the overall asset supply \tilde{S} , 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 critique, according to which \tilde{S} is difficult to identify. The only assumption needed is that the market benchmark (of all actively traded assets) is spanned by all stocks that enter into the covariance matrix Σ .

In the general case when $\lambda < 1$, arbitrageurs take positions in order to benefit from their knowledge about the expected premium change $\Sigma(w^n - w^o)$ in their trading against liquidity suppliers. For arbitrageurs, optimization in the mean–variance space consists of a portfolio choice that linearly combines a return-seeking position with a risk-hedging position. The return-seeking position is achieved by a portfolio that is proportional to the premium or beta change—namely, $\Sigma(w^n - w^o)$. To understand the risk-hedging position, it is useful to calculate the absolute portfolio risk of the return-seeking position as $(w^n - w^o)' \Sigma \Sigma(w^n - w^o)$. 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 with positive marginal arbitrage risk contributions and increases weights in stocks with negative marginal risk contributions. An optimal arbitrage portfolio combines the return-seeking and the risk-hedging 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. The MSCI index revision allows for a straightforward test of these parameter restrictions.

Next, I discuss the asset price behavior after buildup of the speculative positions. As the moment t_L approaches when liquidity suppliers learn about the index change, speculators continuously reduce their hedging positions until they are fully liquidated. The gradual liquidation of hedging positions reverses the price effect that came with their acquisition at time t_A . This price reversal is captured in the following proposition:

Proposition 2 (Excess returns after the speculative position buildup). Over the interval $[t_A, t_L)$, speculators liquidate their hedging positions before liquidity suppliers learn at time t_L about the index change from old index weights w^o to new index weights w^n . The corresponding excess return is positively proportional to the arbitrage risk $\Sigma \Sigma(w^n - w^o)$, where Σ denotes the covariance matrix of asset returns. Formally, the following linear approximation is obtained:

$$\Delta r_{[t_A, t_L]} \approx \beta_2 \Sigma \Sigma(w^n - w^o), \quad (12)$$

with $\beta_2 = -\beta_1 = (1 - \lambda)\gamma \left(\frac{\rho}{\lambda}\right)^2 \vartheta(T - t_u)(t_L - t_A) > 0$.

Proof. See the Appendix.

Proposition 2 characterizes the excess return due to gradual liquidation of the speculators' hedging position. Stocks with high marginal arbitrage risk $[\Sigma \Sigma(w^n - w^o)]_j$ are initially sold short at time t_A . This creates the negative price effect that is captured by the hedging term in Proposition 1. Thereafter, these stocks are gradually bought back until complete liquidation of the hedging position at time t_L . The excess return from the acquisition of the hedging positions exactly offsets the price effect of their liquidation, since $\beta_2 = -\beta_1 > 0$.

The full price dynamics is illustrated in Figure 1. The bold (red) line represents the price $p_t^j - \mathcal{E}(p_T^j)$ net of the expected liquidation value for a stock j with high arbitrage risk. The price effect at time t_A can be decomposed into the return-seeking component given by $\alpha_1 \Sigma(w^n - w^o)$ and the risk-hedging component given by $\beta_1 \Sigma \Sigma(w^n - w^o)$. The latter effect is reversed owing to gradual liquidation of the hedging position over the interval $[t_A, t_L)$. This leads to the V-shaped pattern for a stock with positive arbitrage risk. The return-seeking component fully anticipates the modified stock beta that discounts the liquidation value over the interval $[t_u, T)$. Figure 1 illustrates the case of a lower discount rate (for a beta decrease) by a lower (red) slope for the present value of the stock's liquidation value relative to the initial (blue) present value line.

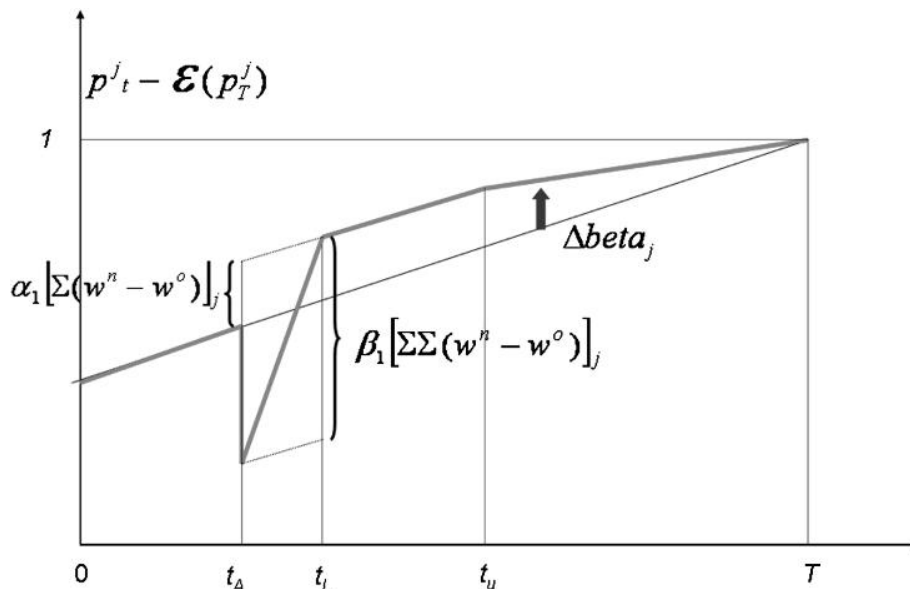


Figure 1

The price dynamics for asset j is depicted net of the expected liquidation value $\mathcal{E}(p_T)$ for the case in which the arbitrage risk $\Sigma \Sigma(w^n - w^o)$ is positive. At time t_A , risk arbitrageurs learn about the demand shock $(w^n - w^o)$, which occurs at time t_u . Liquidity suppliers learn about the demand shock at time $t_L > t_A$.

2.3 Market integration versus segmentation

An important issue in international finance is the degree of integration of different national stock markets. Are asset prices determined locally or globally? Under the hypothesis of national market segmentation, the n assets may be partitioned into m national stock markets according to their primary listing. Speculation 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 that ignores cross-country correlations between stocks listed in different countries by setting those to zero. Formally, the restricted (local) covariance matrix is defined as

$$(\Sigma^L)_{ij} = \begin{cases} 0 & \text{if stocks } i \text{ and } j \text{ are listed in different countries,} \\ (\Sigma^G)_{ij} & \text{otherwise;} \end{cases} \quad (13)$$

here, 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.

Proposition 3 (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 excess return of the speculative position buildup can be decomposed into its local and international components as

$$\begin{aligned} \Delta r_{t_A} &\approx \alpha_1^L \Sigma^L (w^n - w^o) + \alpha_1^{Int} \Sigma^{Int} (w^n - w^o) \\ &+ \beta_1^L \Sigma^L \Sigma^L (w^n - w^o) + \beta_1^{Int} \Sigma \Sigma^{Int} (w^n - w^o) \end{aligned} \quad (14)$$

and the excess return due to liquidated hedging positions as

$$\Delta r_{[t_A, t_L]} \approx \beta_2^L \Sigma^L \Sigma^L (w^n - w^o) + \beta_2^{Int} \Sigma \Sigma^{Int} (w^n - w^o), \quad (15)$$

with

- (i) $\alpha_1^L = \alpha_1^{Int} > 0$, and $\beta_1^L = \beta_1^{Int} < 0$, and $\beta_2^L = \beta_2^{Int} > 0$,
- (ii) $\alpha_1^L > \alpha_1^{Int} = 0$, and $\beta_1^L < \beta_1^{Int} = 0$, and $\beta_2^L > \beta_2^{Int} = 0$

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

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

The intuition behind the test of market integration is straightforward. If stocks share a common market benchmark, then any contribution to a benchmark change should, *ceteris paribus*, not depend on whether this contribution comes from the home market or a foreign market. Market integration is characterized by the symmetrical roles played by home and foreign stocks in defining the market benchmark and also its change. It is this symmetrical benchmark role of all globally priced equity that is tested by decomposing the global covariance matrix.

This interpretation can be illustrated further by the following example. Assume that the stock price of Microsoft (stock m) covaries equally strongly with the stock return of General Electric (stock g) and the Italian company Fiat (stock f) and that both GE and Fiat are up-weighted in the MSCI index by the same amount; hence, $(w^n - w^o)_g = (w^n - w^o)_f > 0$. Under market integration, the index weight increase of both GE and Fiat should produce quantitatively the same long-run effect on the beta and stock price of Microsoft because $\Sigma_{mg}^G (w^n - w^o)_g = \Sigma_{mf}^G (w^n - w^o)_f$. This equality of the cross-border pricing effects is tested by separating the GE element $\Sigma_{mg}^G (w^n - w^o)_g$ as part of the local premium change $\Sigma_{m\bullet}^L (w^n - w^o)$ from the Fiat element $\Sigma_{mf}^G (w^n - w^o)_f$ as part of the international premium change $\Sigma_{m\bullet}^{Int} (w^n - w^o)$. 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 Microsoft risk premium, then its change should be without consequence for the Microsoft stock price and so $\alpha_1^{Int} = 0$. 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 arbitrated by U.S. investors, Italian stocks only by Italian investors, and so forth. In this case, the submatrix $\Sigma^L \Sigma^L$ is sufficient to characterize all arbitrage risk; we therefore expect a zero contribution from the international arbitrage risk component $\Sigma \Sigma^{Int}$, or $\beta^{Int} = 0$. However, the latter 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 this case, stock markets are integrated with respect to arbitrage behavior.

The described procedure is easily adapted to market segmentation tests along data dimensions other than the national stock market listing. I only need to decompose the matrices Σ^G and $\Sigma^G \Sigma^G$ differently to obtain analogous tests of market segmentation between, for example, developed market and

emerging market stocks, or only between developed market and illiquid emerging market stocks.

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, with or without dividends reinvested. MSCI's global equity indices have become the most widely used international equity benchmarks by institutional investors. By the year 2000, nearly 2,000 organizations worldwide were using the MSCI international equity benchmarks. Over \$3 trillion of investments were benchmarked against these indices worldwide, and approximately \$300–350 billion were directly indexed.¹⁴ The index with the largest international coverage is the MSCI All Country World Index (ACWI), which includes 50 developed and emerging equity markets. This broad index is the focus of the empirical work reported here. MSCI reviews the index composition at regular intervals in order to maintain a broad and fair market representation.¹⁵ In 2000, however, MSCI initiated an index review of exceptional scope as described in the following section.

3.1 Public announcement and the event windows

In February 2000, MSCI started to review its weighting policy and 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 investability of many stocks. Free-float weights were adopted by MSCI's competitor Dow Jones on September 18, 2000. The next day, MSCI published a consultative paper on possible changes and sought comments from its clients. The consultation process between MSCI and the investment industry proceeded throughout November 2000. It is thus more than 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, 2000, 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-float weights was highly

¹⁴ See the investment newsletter "Spotlight on: Throwing Weights Around," Hewitt Investment Group, December 2000.

¹⁵ 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 reassesses systematically the various dimensions of the equity universe for all countries. Second, there are index reviews at the end of February, August, and November, in which other significant market events are accounted for (e.g., large market transactions affecting strategic shareholders, exercise of options, share repurchases). Third, ongoing event-related changes such as mergers and acquisitions, bankruptcies, or spinoffs are implemented as they occur.

probable. The second announcement, on December 10, 2000, provided the official confirmation that MSCI would adopt free-float weights. MSCI also communicated the timetable for implementation of the index change in two steps as well as the new target for market representation: an increase to 85% from the previous level of 60%. The equity index would adjust 50% toward 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 previously circulated consultative paper. The target level of 85% was somewhat higher (by 5%), and the implementation timetable somewhat longer, than industry observers had expected.¹⁶

It is most plausible that arbitrageurs acquired their speculative positions during the month of November, in parallel with MSCI's consultation process. Thus, speculative positions are likely to have been built up even before the first announcement on December 1. Since the exact beginning of the speculative activity is difficult to date, a variety of different event windows are proposed, all of which extend until the market closure on December 1. These windows capture the "position buildup event" and cover alternatively a period of 5, 10, 15, or 20 trading days. Their event returns should capture the excess return Δr_{t_A} associated with time t_A in the model.

Liquidity providers and the market as a whole may have revised their stock valuation much later than December 1, 2000. After all, knowing exactly which stocks would be up- or down-weighted required considerable equity research into the ownership structure of more than 2,300 stocks. Following the weekend of December 2 and 3, 2000, the financial market reopened on December 4, and the market closure on this date is chosen as the beginning of a second event window. It captures the excess returns $\Delta r_{[t_A, t_L]}$ predicted in Proposition 2 that are associated with the liquidation of hedging positions as the moment t_L of symmetric information approaches. Here again, different event windows are selected: Starting on December 4, 2000, they extend over the following 3, 5, or 7 trading days. Different window lengths should aid in assessing the robustness of the findings.¹⁷

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 rather than market capitalization. Second, the market representation is enhanced in the new index. Both changes entail rule-based weight changes that do not involve subjective judgments about the growth prospects of a stock.

¹⁶ See again the investment newsletter "Spotlight on: Throwing Weights Around," Hewitt Investment Group, December 2000.

¹⁷ The reported regression outcomes also feature a certain robustness with respect to the end date of the first event window and the beginning of the second event window. For example, extending the "position buildup event" until December 4 or starting the "hedge liquidation event" on December 1 has no incidence on the qualitative results.

This implies that a speculator's correct anticipation of the rule change allows her to calculate the individual stock weight changes.

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 commonly result from "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 or 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; non-free-floating shares include those held by governments, companies, banks (excluding 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% within each industry and country, compared to the 60% share based on market capitalization in the old index.

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 2,077 stocks. The new index methodology led to the addition of 489 new stocks and the removal of 298 stocks. The total number of stocks belonging either to the old or new index is therefore 2,566. 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 Argentina and Turkey: Both countries had currency pegs (to the U.S. dollar) that were coming under increasing pressure, so their stocks could be subject to speculative considerations outside our model framework.¹⁸ The analysis also requires two years of historic price data to compute covariance matrices with all other index stocks. Datastream is used as the source for all price data. For 31 stock codes, no company information was found. Another 182 stocks had fewer than 80 weekly return observations for the two-year period prior to the index change and are thus also discarded. This reduces the data sample from 2,566 to 2,291 stocks, of which 396 are included and 265 excluded in the index revision. Two robustness checks are undertaken with respect to these sample selection criteria. First, including stocks from Argentina and Turkey increases the sample to 2,349 stocks. Second, an inclusion threshold of only 40 weekly return observations expands the sample further to 2,414 stocks. Both expanded samples produced qualitatively similar results to those reported here. Companies with missing observations tend to have small weights and

¹⁸ The Turkish lira lost 28% of its value upon floating on February 22, 2001. A first devaluation of the Argentine peso by 29% occurred in January 2002, and the subsequent abandoning of its peg implied another 75% value loss with respect to the U.S. dollar in a matter of months. I also note that the risk of exchange rate controls for both countries makes their stocks problematic investments for global arbitrage strategies.

Table 1
Revision of the global MSCI equity index by country

Country	(1) MSCI Stocks	(2) Sample Stocks	(3) New Weight	(4) Old Weight	(5) Mean(Δv^j)	(6) SD(Δv^j)	(7) EM Dummy
Argentina	17	0	0.03	0.07	-	-	1
Australia	71	66	1.28	1.29	0.395	0.939	0
Austria	17	16	0.04	0.09	-0.702	1.021	0
Belgium	18	16	0.28	0.39	-0.308	0.991	0
Brazil	50	47	0.30	0.49	-0.638	1.041	1
Canada	89	80	1.93	1.97	0.195	1.025	0
Chile	30	30	0.07	0.18	-0.777	0.941	1
China	43	37	0.26	0.35	0.091	1.215	1
Colombia	9	8	0.00	0.01	-1.184	0.560	1
Czech Republic	6	6	0.01	0.03	-0.958	0.149	1
Denmark	27	24	0.28	0.40	0.098	1.211	0
Egypt	14	11	0.01	0.01	0.033	0.000	1
Finland	30	27	1.00	0.98	-0.598	1.116	0
France	63	59	3.77	4.93	-0.279	1.03	0
Germany	59	51	2.76	3.71	-0.276	1.013	0
Greece	24	21	0.16	0.16	0.033	0.000	0
Hong Kong	34	32	0.66	0.93	-0.511	1.137	1
Hungary	13	13	0.03	0.04	-1.079	0.913	1
India	76	74	0.12	0.35	-1.037	0.943	1
Indonesia	32	32	0.02	0.03	-1.171	1.039	1
Ireland	16	15	0.34	0.30	0.044	0.942	0
Israel	52	49	0.16	0.24	-0.765	1.407	1
Italy	47	41	1.38	1.99	-0.436	1.107	0
Japan	348	333	9.38	10.71	0.035	1.068	0
Korea	82	69	0.45	0.57	-0.119	1.102	0
Luxembourg	3	3	0.01	0.05	-1.093	0.478	0

(continued)

Table 1
Continued

Country	(1) MSCI Stocks	(2) Sample Stocks	(3) New Weight	(4) Old Weight	(5) Mean(Δv^j)	(6) SD(Δv^j)	(7) EM Dummy
Malaysia	86	84	0.12	0.29	-0.696	1.186	1
Mexico	27	25	0.3	0.59	-0.674	0.938	1
Morocco	12	11	0.01	0.01	0.033	0.000	1
Netherlands	26	24	2.17	2.42	-0.045	0.665	0
New Zealand	15	10	0.04	0.05	0.441	1.446	0
Norway	24	22	0.16	0.20	0.007	0.834	0
Pakistan	18	18	0.00	0.01	-1.339	0.706	1
Peru	11	10	0.01	0.02	-0.294	1.344	1
Philippines	21	20	0.02	0.05	-0.901	1.176	1
Poland	22	19	0.04	0.06	-0.404	0.945	1
Portugal	11	10	0.15	0.23	-0.728	0.492	0
Russia	12	9	0.08	0.14	-0.550	1.178	1
Singapore	40	34	0.26	0.38	-0.221	1.167	1
South Africa	48	45	0.47	0.55	-0.378	0.922	1
Spain	34	31	1.21	1.38	-0.830	0.964	0
Sri Lanka	8	8	0.00	0.00	-0.753	0.589	1
Sweden	37	35	0.84	1.00	-0.022	0.647	0
Switzerland	43	38	2.87	2.93	-0.105	1.143	0
Taiwan	96	89	0.44	0.69	0.111	1.304	1
Thailand	39	37	0.04	0.08	-0.926	1.176	1
Turkey	45	0	0.05	0.12	-	-	1
United Kingdom	140	133	10.33	9.26	0.369	0.924	0
United States	443	414	55.12	48.88	0.390	1.108	0
Venezuela	7	5	0.01	0.02	-0.598	1.470	1
Without company data	31	0	0.5	0.38	-	-	-
Total	2,996	2,291	100.00	100.00	-	-	-

Reported are summary statistics by country on the (1) total number of stocks affected by MSCI index revision; (2) total number of sample stocks with complete historic price data; (3) new; and (4) old country weights in percent. Also provided for the sample stocks are the (5) mean; and (6) standard deviation (SD) of the percentage weight change $\Delta v^j = 2(w^n - w^o)_j / (w^n + w^o)_j$ within a country. Column (7) marks countries with emerging market (EM) stocks.

small absolute weight changes so that their influence on the market benchmark change is negligible.

Columns (3) and (4) of Table 1 list (based on the sum of new and old stock weights) the new and old aggregate country weights, respectively. As the new stock weights, I use new (provisional) weights announced on May 31, 2001. These early (provisional) weights should be close to what the market anticipated about the new index weights in November 2000. The largest contribution to the new MSCI index comes from U.S. stocks (55.12%), followed by the United Kingdom (10.33%), and Japan (9.38%). The most dramatic country weight change is seen in the United States (a 6.24% absolute weight increase), followed by the United Kingdom (a 1.07% increase). Both countries also account for 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

$$\Delta v_j = \frac{w_j^n - w_j^o}{\frac{1}{2}(w_j^n + w_j^o)}, \quad (16)$$

where w_j^o and w_j^n represent (respectively) the old and new index weights of stock j . The percentage weight change is bounded above by +2 for newly included stocks and below by -2 for deleted stocks. Columns (5) and (6) of Table 1 report the mean and the standard deviation (SD) of the percentage weight change Δv_j by country. The largest average stock weight increase is experienced by stocks in New Zealand (44.1%), the United States (39.0%), and the United Kingdom (36.9%). Figure 2 plots the percentage weight change of individual stocks against their initial weight (in logs) for non-U.S. stocks as well as U.S. stocks. Because of 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%.

3.3 Risk premium changes and marginal arbitrage risk

In order to determine the premium change and the arbitrage risk, I need to estimate the covariance matrix Σ of all stock returns. To proxy for the (expected) covariance matrix, I simply use the historical covariance based on two years of return data prior to the event. The estimation window for the covariance covers the period July 1, 1998, to July 1, 2000; this is sufficiently removed from the first announcement on December 1, 2000, to be unaffected by the event itself. The covariance estimation for the stock returns is based on weekly data. Because stock prices are sampled around the world, daily sampling may pose inference problems due to asynchronous return measurement. Weekly return sampling seems more robust to this problem, which justifies the use of weekly data. On a more general level, using historical data certainly represents an

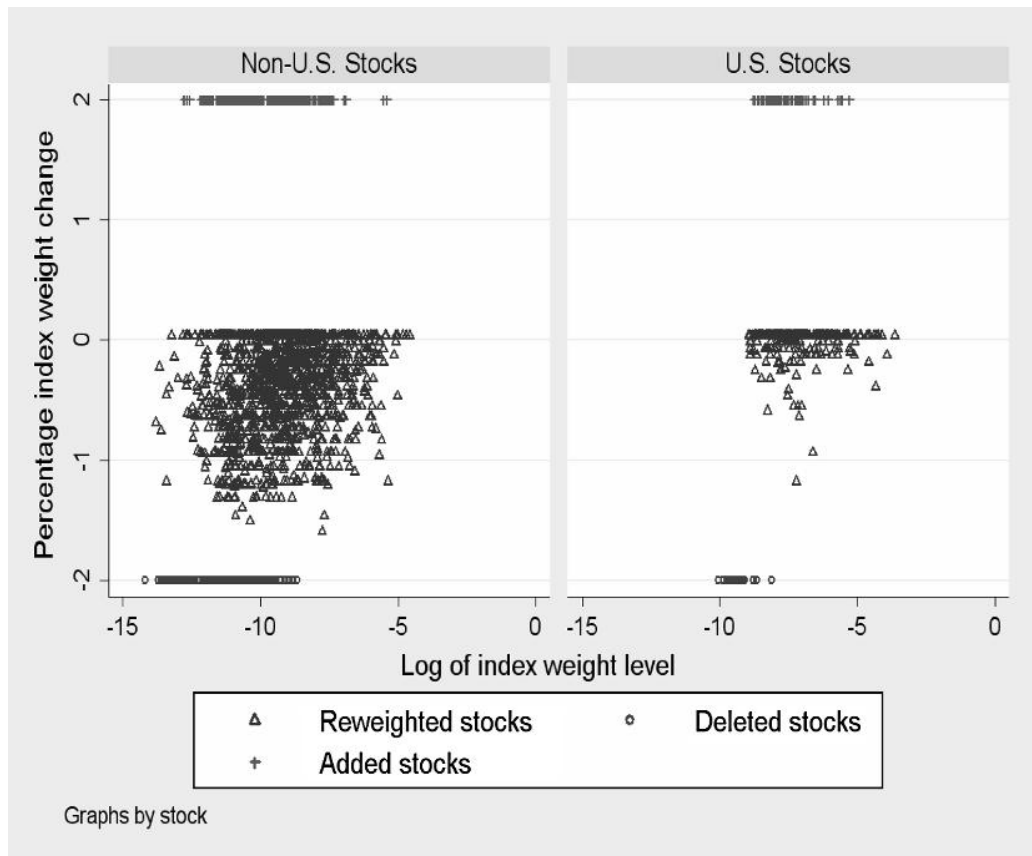


Figure 2

The percentage weight change for non-U.S. and 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). The graph distinguishes between reweighted stocks, stocks added to the index, and stocks deleted from the index.

imperfect measure of the forward-looking covariance, but it is also the method most likely used by arbitrageurs to determine the optimal arbitrage strategy and the ex ante risk of their portfolio position. It is important to emphasize 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 2,291 different return sequences. The estimation quality is therefore comparable to the standard beta estimation.

The most important 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 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 is 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 that 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 is 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.

Equation (3) in Section 1.1 relates the premium changes $\Sigma(w^n - w^o)$ to corresponding beta changes. An economic interpretation of the magnitude of beta changes depends on the parameter ϑ , which represents the ratio of MSCI index capital to market benchmark capital and requires additional discussion. MSCI index capital may be defined narrowly as the index-tracking capital or, more broadly, as the capital that has the MSCI index as its performance benchmark and therefore maintains portfolio weights relatively close to the index. MSCI's own estimate of capital that is benchmarked to the index is more than \$3 trillion for the year 2000. The amount of capital in the market portfolio is more difficult to evaluate. An asset pricing model (like the CAPM model) should hold for unconstrained investors who are continuously optimizing their risk–return trade-off. A large share of institutional equity investments can be regarded as passively invested because it closely tracks performance benchmarks. Recall that passively invested capital does not count toward capital in the market portfolio; household capital, on the other hand, may be subject to behavioral investment biases or infrequent portfolio adjustment. Control-related equity holdings also play an important role, especially outside the United States (Kho, Stulz, and Warnock 2009), and should further reduce the market benchmark capital. Generally, the smaller the amount of market benchmark capital, the larger the parameter ϑ and the more important becomes the beta change of the index revision.

Some simple calculations show which particular parameter assumptions produce sizable beta changes and return effects. Assume, for example, that 10% of global equity capital stock (at \$36 trillion in 2000) counts as market benchmark capital. An MSCI index capital stock of \$3 trillion then yields a parameter $\vartheta = 0.83$. The weekly market return variance of the global index is estimated as $\sigma_m^2 = w^o \Sigma w^o = 0.936$. The standard deviation (SD) for the term $[\Sigma^G(w^n - w^o)]_j$ is given by 0.049 (see Table 2, Panel A), which implies 0.044 ($= 0.83 \times 0.049/0.936$) for the SD of global beta changes. The average stock price effect can then be evaluated based on the discount rate variation for a growing perpetuity. If one assumes a 3% risk-free rate and an equity premium of 5%, then a beta decrease of 0.044 (relative to an initial beta of 1)

Table 2
Summary statistics on stock premium changes and marginal arbitrage risk

Sample	Variable	Obs.	Mean	SD	Min.	Max.
<i>Panel A: Change in Risk Premium under Global Pricing</i>						
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	$[\Sigma^G(w^n - w^o)]_j$	1,877	-0.009	0.036	-0.173	0.219
<i>Panel B: Change in Risk Premium under Local Pricing</i>						
All Stocks	$[\Sigma^L(w^n - w^o)]_j$	2,291	0.016	0.057	-0.074	0.343
Added and Deleted Stocks	$[\Sigma^L(w^n - w^o)]_j$	661	0.027	0.072	-0.072	0.343
U.S. Stocks	$[\Sigma^L(w^n - w^o)]_j$	414	0.115	0.074	-0.073	0.343
Non-U.S. Stocks	$[\Sigma^L(w^n - w^o)]_j$	1,877	-0.006	0.010	-0.051	0.024
<i>Panel C: Risk Contribution to Global Arbitrage Portfolio</i>						
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	$[\Sigma^G \Sigma^G(w^n - w^o)]_j$	1,877	9.42	46.62	-179.46	302.09
<i>Panel D: Risk Contribution to Local Arbitrage Portfolio</i>						
All Stocks	$[\Sigma^L \Sigma^L(w^n - w^o)]_j$	2,291	19.67	57.17	-70.62	390.30
Added and Deleted Stocks	$[\Sigma^L \Sigma^L(w^n - w^o)]_j$	661	29.42	74.18	-70.62	390.30
U.S. Stocks	$[\Sigma^L \Sigma^L(w^n - w^o)]_j$	414	119.09	77.15	-70.62	390.30
Non-U.S. Stocks	$[\Sigma^L \Sigma^L(w^n - w^o)]_j$	1,877	-2.26	4.05	-20.13	7.16
<i>Panel E: Market Integration across Emerging and Developed Markets</i>						
All Stocks	$[\Sigma^H(w^n - w^o)]_j$	2,291	0.008	0.051	-0.145	0.29
All Stocks	$[\Sigma^{CH}(w^n - w^o)]_j$	2,291	0.002	0.029	-0.156	0.20
All Stocks	$[\Sigma^H \Sigma^H(w^n - w^o)]_j$	2,291	22.79	60.53	-131.42	358.82
All Stocks	$[\Sigma \Sigma^{CH}(w^n - w^o)]_j$	2,291	3.09	39.64	-177.27	279.32
<i>Panel F: Market Integration by Cross Listing and Liquidity Characteristics</i>						
All Stocks	$[\Sigma^{List+}(w^n - w^o)]_j$	2,291	-0.000	0.016	-0.062	0.200
All Stocks	$[\Sigma^{List-}(w^n - w^o)]_j$	2,291	-0.002	0.024	-0.152	0.187
All Stocks	$[\Sigma^{Liq+}(w^n - w^o)]_j$	2,291	-0.002	0.024	-0.065	0.200
All Stocks	$[\Sigma^{Liq-}(w^n - w^o)]_j$	2,291	-0.001	0.016	-0.156	0.182
<i>Panel G: Control Variables</i>						
All Stocks	PP^j	2,291	-0.129	1.149	-2.000	2.000
All Stocks	Liq^j	2,291	-0.132	0.126	-1.321	0.000
All Stocks	$PP^j \times Liq^j$	2,291	0.044	0.251	-2.036	2.344

Panels A and B report summary statistics on global and local stock risk premium changes, respectively; Panels C and D provide the same statistics on global and local arbitrage risk contributions. The global covariance matrix is denoted by Σ^G . In the local covariance matrix Σ^L , matrix elements are set to zero for stocks in different national markets. In Panel E, the matrix Σ^H captures only covariances within the hemisphere of developed market stocks and within the hemisphere of emerging market stocks; it is zero pairs of developed and emerging market stocks. The (cross-hemisphere) covariances between developed and emerging market stocks are represented by the matrix $\Sigma^{CH} = \Sigma^G - \Sigma^H$. In Panel F, the latter covariance matrix Σ^{CH} is further decomposed into (i) stock pairs with and without a cross-listing for the emerging market stock, $\Sigma^{CH} = \Sigma^{List+} + \Sigma^{List-}$; and (ii) stock pairs for which the emerging market stock has above- or below-median liquidity, $\Sigma^{CH} = \Sigma^{Liq+} + \Sigma^{Liq-}$. All covariance matrices are estimated for two 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. Panel G reports summary statistics of the price pressure proxy defined as the percentage stock weight change $PP^j = \Delta v^j = 2(w^n - w^o)_j / (w^n + w^o)_j$ and the stock liquidity proxy defined as $Liq^j = \ln(1 = ZR^j)$, where ZR^j denotes the percentage of zero daily returns over a prior two-year period.

produces a 22-basis-point ($= 0.05 \times 0.044$) drop in the discount factor and a 5.82% ($= -1 + (0.08 - 0.04)/(0.0778 - 0.04)$) value increase for a cash flow perpetuity growing at 4%. This is economically significant and comes close to the observed return dispersion.¹⁹ These calculations show that the index change is likely to produce a sizable return effect if the MSCI index capital and the market benchmark capital are of similar magnitude. Conversely, observing significant return effects suggests that the amount of market benchmark capital may not exceed index capital by an order of magnitude.

Next I discuss the difference between global and local premium changes. A graphical representation of the distribution of the global and local risk premium change is provided in Figure 3, which 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. Compared with the global covariance matrix Σ^G , the local covariance matrix Σ^L features (by construction) many zero elements, and this tends to generate less dispersion in the local than in the global premium change. The dispersion of local premium changes is especially small for stocks from countries with a minor representation in

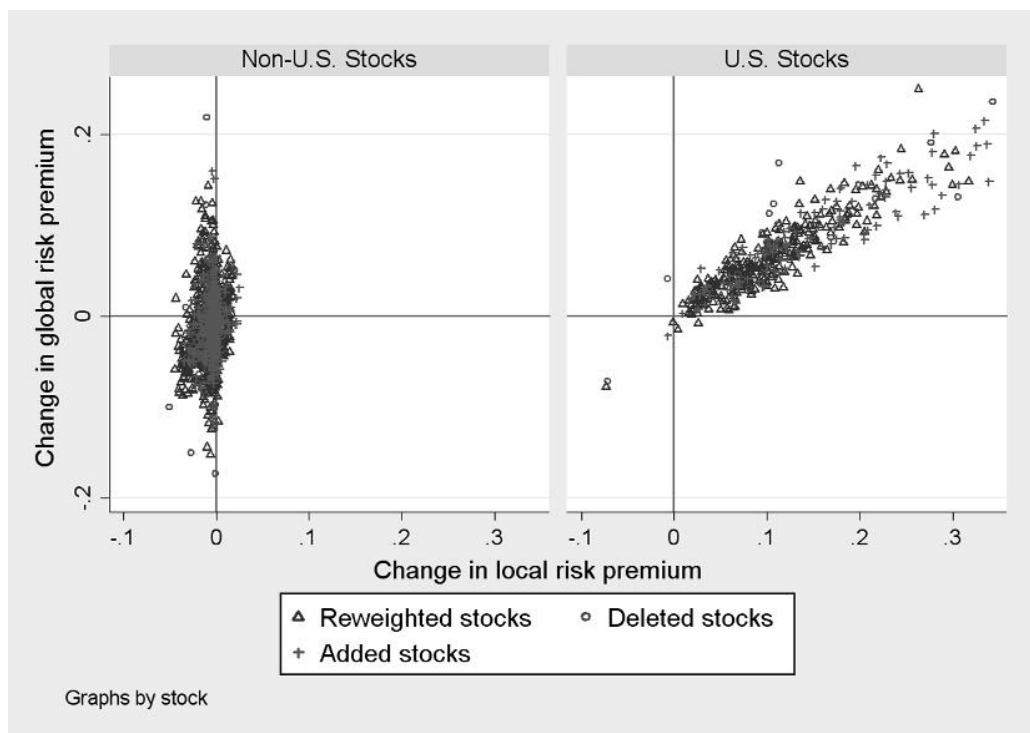


Figure 3 For non-U.S. and U.S. stocks, the risk premium change $[\Sigma^G(w^n - w^o)]_j$ in stocks j under global asset pricing (market integration) is plotted against the risk premium change $[\Sigma^L(w^n - w^o)]_j$ under local asset pricing (market segmentation). The graph distinguishes reweighted stocks, stock added to the index, and stocks deleted from the index.

¹⁹ The estimated price effect of a 1 SD change in $[\Sigma(w^n - w^o)]_j$ over the fifteen-day window in Table 3 is approximately 5% ($\approx 101.7 \times 0.049$). The latter estimate corresponds to $\vartheta = 0.75$.

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 corresponds to the correlation of the local and global beta changes and can be calculated as 0.234. This low correlation allows for sufficient discriminatory power between local and global asset pricing effects.

For U.S. stocks, the premium changes behave quite differently. The local equity premium change for U.S. stocks shows a large SD of 0.074. The local premium change here is typically only slightly smaller than the global premium change, as seen in Figure 3. Most U.S. stocks are situated just below the 45-degree line. The large number of U.S. stocks in the MSCI index explains why the corresponding rows in the global and local covariance matrices differ less for U.S. stocks than for stocks from other countries: Fewer cross-country covariances are set to zero. As a consequence, local and global premium and beta changes show a high correlation (0.912) for U.S. stocks, which makes them 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.

The distribution of local and global marginal arbitrage risk is 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, whereas this correlation is 0.987 for U.S. stocks, indicating strong collinearity. Meaningful inference about global versus local arbitrage risk is therefore problematic for the sample of only U.S. stocks.

4. Event Evidence

Evidence on the price impact of the speculative position buildup is presented in Section 4.1. Once the speculators have acquired their desired positions, they liquidate their hedging positions over the subsequent hedge liquidation period examined in Section 4.2. The evidence on global versus local market pricing is discussed in Section 4.3, followed by evidence on market segmentation between developed and emerging market stocks in Section 4.4. Section 4.5 presents evidence on the global market integration of emerging market stocks sorted by cross-listing and liquidity characteristics.

4.1 Price effects of the speculative position buildup

The global scale of the MSCI index rebalancing provides an extremely large sample of stocks that experienced a weight change. The sample contains 2,291 stocks, along with a continuous two-year price history that is needed to calculate the global covariance matrix Σ^G . The statistical inference is based

on a cross-sectional analysis in which dollar returns Δr^j (defined as log price differences $\ln P_t^j - \ln P_{t-\Delta t}^j$) in stock j over the entire event window are regressed on a constant c , the stock's risk premium change $[\Sigma^G(w^n - w^o)]_j$, and its corresponding marginal arbitrage risk $[\Sigma^G \Sigma^G(w^n - w^o)]_j$. Formally,

$$\Delta r_{tA}^j = c + \alpha_1[\Sigma^G(w^n - w^o)]_j + \beta_1[\Sigma^G \Sigma^G(w^n - w^o)]_j + \mu_j, \quad (17)$$

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 windows are considered: 5, 10, 15, or 20 trading days prior to December 1, 2000. Panel A of Table 3 shows the regression results for the full sample of 2,291 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, which represents the special case $\lambda = 1$ where all market participants are equally informed arbitrageurs and there is no liquidity supply. This restrictive specification is strongly rejected by the data. The estimated coefficient α_1 is negative, whereas theory predicts a positive coefficient. The Greenwood model is rejected for each of the four event windows. However, 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% for two stocks with a relative change in their risk premium of 1 SD 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 of 1 SD (or 61.63) in a particular stock induces smaller (or short) positions and therefore a 6.1% decrease in the 10-day pre-announcement return. At 0.119, the adjusted R^2 of the full specification is highest for the 20-day event window and more than two times higher than under the restrictive specification. The estimated coefficients for the full specification increase with the window size—as 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 show qualitatively similar results. In each case and for every window size, the hypothesis $\beta_1 = 0$ is strongly rejected. As with the entire sample and in line with the theoretical model—in the full specification, the coefficient α_1 for the risk premium change is significantly positive and the coefficient β_1 for the arbitrage risk significantly negative. The adjusted R^2 is higher in Panel B, suggesting a better model fit for stocks with the most dramatic weight changes. For the 20-day window in Panel B, an adjusted R^2 of 0.18 is found.

Table 3
Price effect of the speculative position buildup

WS	c	$[t]$	α_1	$[t]$	β_1	$[t]$	R^2
<i>Panel A: Position Buildup Event (All Stocks, N = 2,291)</i>							
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
<i>Panel B: Position Buildup Event (Only Added and Deleted Stocks, N = 661)</i>							
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
<i>Panel C: Position Buildup Event (Only Non-U.S. Stocks, N = 1,877)</i>							
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

To characterize the price effect of the speculative position buildup, the cumulative event returns Δr_{tA}^j (denominated in dollars and expressed in percentage points) over different even windows (WS = window size) is regressed on a constant, the change in the risk premium $[\Sigma^G(w^n - w^o)]_j$, and the arbitrage risk $[\Sigma^G \Sigma^G(w^n - w^o)]_j$ of each stock j . Formally,

$$\Delta r_{tA}^j = c + \alpha_1[\Sigma^G(w^n - w^o)]_j + \beta_1[\Sigma^G \Sigma^G(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for two years of weekly dollar stock returns for the period of July 1, 1998, to July 1, 2000. The event window size is chosen in turn 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 brackets.

To verify that the measured effects are not due primarily to country-level variation in the independent variables (instead of stock-level variation), all regressions are replicated with country fixed effects; this still yields qualitatively similar results. Another robustness check consists in excluding all *technology*,

media, and telecommunication companies from the regression. Such stocks might have been characterized by extreme within-group correlations around their valuation peak in March 2000, but qualitatively similar results again persist. Overall, returns for the position buildup event provide strong empirical support for the generalized arbitrage model. The estimated effects are also economically significant.

4.2 Price effects of liquidating hedging positions

According to Proposition 2, the return effect of the hedge positions is reversed after their buildup. Arbitrageurs gradually liquidate their hedging positions as date t_L approaches, and prices reflect all new information. I assume that liquidation of the hedging positions starts at the end of the position buildup window. Three alternative windows for the *hedge liquidation event* are defined: These event windows all start on the market closure on December 4, 2000, and cover returns over three, five, or seven trading days. The five-day window extends until the market closure on December 11, 2000—the first trading day after MSCI’s decision becomes public.

From Proposition 2, the preferred cross-sectional specification for the post-announcement return effect follows as

$$\Delta r_{[t_A, t_L]}^j = c + \beta_2 [\Sigma^G \Sigma^G (w^n - w^o)]_j + \mu_j, \quad (18)$$

where $\beta_2 > 0$ is expected for the general model with $\lambda < 1$. 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 that $\beta_2 = 0$ under $\lambda = 1$.

Panel A of Table 4 reports regression results for the full sample. The reported t -statistics are again robust to error clustering at the country level. The coefficient β_2 is significant at the 1% level in all specifications, in all samples, and for all three event windows. The adjusted R^2 reaches 0.138 for the five-day event window; at 0.175, it is even higher for the sample of added and deleted stocks reported in Panel B. Overall, the post-announcement return pattern provides additional support for favoring the generalized model of risk arbitrage over the nested Greenwood framework.

4.3 Global versus local asset pricing

Arbitrage strategies could employ all MSCI stocks or only a subset of stocks in the local (i.e., national) market. The investor mandate might constrain some fund managers from investing 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

$$[\Sigma^{Int} (w^n - w^o)]_j = [\Sigma^G (w^n - w^o)]_j - [\Sigma^L (w^n - w^o)]_j \quad (19)$$

Table 4
Price effect of liquidating hedging positions

WS	c	$[t]$	β_2	$[t]$	R^2
<i>Panel A: Hedge Liquidation Event (All Stocks, N = 2,291)</i>					
3	0.67	[1.30]	0.023	[3.90]	0.037
5	1.90	[2.90]	0.047	[6.78]	0.138
7	1.40	[2.17]	0.033	[4.56]	0.069
<i>Panel B: Hedge Liquidation Event (Only Added and Deleted Stocks, N = 661)</i>					
3	0.74	[1.79]	0.023	[7.08]	0.067
5	2.06	[3.51]	0.054	[6.26]	0.175
7	1.52	[2.22]	0.034	[6.81]	0.075
<i>Panel C: Hedge Liquidation Event (Only Non-U.S. Stocks, N = 1,877)</i>					
3	1.00	[2.14]	0.032	[4.06]	0.089
5	2.37	[4.65]	0.044	[3.76]	0.093
7	1.76	[2.84]	0.041	[3.30]	0.070

To characterize the price effect of liquidating hedging positions, the cumulative equity returns $\Delta r_{[t_A, t_L]}^j$ (denominated in dollars and expressed in percentage points) for different event windows (WS = window size) are regressed on a constant and the arbitrage risk $[\Sigma^G \Sigma^G (w^n - w^o)]_j$ of each stock j . Formally,

$$\Delta r_{[t_A, t_L]}^j = c + \beta_2 [\Sigma^G \Sigma^G (w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for two years of weekly dollar stock returns for the period of July 1, 1998, to July 1, 2000. The event window size is chosen in turn 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 brackets.

and the incremental international *marginal arbitrage risk* as

$$[\Sigma \Sigma^{Int} (w^n - w^o)]_j = [\Sigma^G \Sigma^G (w^n - w^o)]_j - [\Sigma^L \Sigma^L (w^n - w^o)]_j, \quad (20)$$

where Σ^G represents the covariance of dollar returns for all 2,291 stocks and Σ^L the equivalent covariance matrix with zeros for stocks in different countries. The statistical inference for the position buildup event is based on the regressions

$$\begin{aligned} \Delta r_1^j = & c + \alpha_1^L [\Sigma^L (w^n - w^o)]_j + \alpha_1^{Int} [\Sigma^{Int} (w^n - w^o)]_j \\ & + \beta_1^L [\Sigma^L \Sigma^L (w^n - w^o)]_j + \beta_1^{Int} [\Sigma \Sigma^{Int} (w^n - w^o)]_j + \mu_j \end{aligned} \quad (21)$$

and for the hedge liquidation event on

$$\Delta r_2^j = c + \beta_2^L [\Sigma^L \Sigma^L (w^n - w^o)]_j + \beta_2^{Int} [\Sigma \Sigma^{Int} (w^n - w^o)]_j + \mu_j, \quad (22)$$

where Δr_1^j and Δr_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 of 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 $\beta^{Int} = 0$ suggests strictly local arbitrage strategies—two conditions that characterize an internationally segmented market.

Table 5 reports regression results for the decomposition of the position buildup event into the local and international return components. In Panel A, the sample consists of all sample stocks. First, the economic magnitude of the return effects can be highlighted. Estimates of $\alpha_1^L = 164.1$ and $\beta_1^L = -0.205$ for the 20-day window imply that a 1 SD change in the stock beta or the arbitrage risk modifies stock prices by 9.35% ($= 164.1 \times 0.057$) and -11.72% ($= -0.205 \times 57.17$), respectively. Second, the incremental effects captured by the coefficients α_1^{Int} and β_1^{Int} are statistically significant for each of the event windows and also have the expected sign. The risk premium change and the marginal arbitrage risk therefore have an important international component. The arbitrage strategies assumed the validity of an international premium change and also engaged in international hedging. Third, the magnitude of the international coefficients is similar to that of the corresponding local coefficients. The last two columns in Table 5 report the significance level for an F -test conjecturing equality of the respective coefficients. Neither the null hypothesis $\alpha_1^L = \alpha_1^{Int}$ nor $\beta_1^L = \beta_1^{Int}$ can be rejected. For example, the local beta change produces a coefficient estimate $\alpha_1^L = 164.1$ for the 20-day return in Panel A, and the complementary international beta change (induced by weight changes in foreign country stocks) shows a coefficient estimate of $\alpha_1^{Int} = 127.0$. The corresponding estimates for the marginal arbitrage risk are $\beta_1^L = -0.205$ and $\beta_1^{Int} = -0.153$, respectively.

As a robustness check, separate results are estimated for all non-U.S. stocks. As discussed previously, U.S. stocks are characterized by a relatively high correlation between local and global risk premium changes and also between local and global marginal arbitrage risk. This makes discrimination between the local and global pricing component more difficult. In contrast, non-U.S. stocks feature a much lower correlation between local and global explanatory variable. But their local premium and local arbitrage risk variation are small, and the coefficients α_1^L and β_1^L are statistically insignificant for some of the event windows in Panel B. 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. Just as for the full sample, the hypothesis of equity market integration cannot be rejected, but the hypothesis of local asset pricing is strongly rejected.

Table 5
Market integration test for speculative position buildup

WS	c	$[t]$	α_1^L	$[t]$	α_1^{Int}	$[t]$	β_1^L	$[t]$	β_1^{Int}	$[t]$	R^2	F-Test $\alpha_1^L = \alpha_1^{Int}$	F-Test $\beta_1^L = \beta_1^{Int}$
<i>Panel A: Position Buildup Event (All Stocks, N = 2,291)</i>													
5	1.48	[3.14]	-6.2	[-0.13]	51.8	[3.20]	-0.017	[-0.36]	-0.066	[-5.05]	0.100	0.298	0.358
10	-0.33	[-0.41]	32.6	[0.88]	89.0	[4.01]	-0.051	[-1.47]	-0.101	[-6.49]	0.091	0.109	0.137
15	-1.97	[-1.52]	136.0	[1.66]	96.7	[3.47]	-0.140	[-2.19]	-0.103	[-4.51]	0.052	0.518	0.534
20	-1.79	[-1.22]	164.1	[3.34]	127.0	[3.25]	-0.205	[-4.23]	-0.153	[-4.55]	0.121	0.474	0.278
<i>Panel B: Position Buildup Event (Non-U.S. Stocks, N = 1, 877)</i>													
5	1.34	[5.05]	100.0	[3.47]	34.4	[3.29]	-0.291	[-3.91]	-0.051	[-6.41]	0.070	0.034	0.002
10	-0.78	[-2.57]	40.5	[1.04]	75.7	[5.33]	-0.076	[-0.75]	-0.091	[-7.96]	0.063	0.413	0.884
15	-2.64	[-6.77]	80.9	[1.74]	82.5	[4.84]	0.005	[0.04]	-0.092	[-6.35]	0.037	0.974	0.439
20	-2.75	[-6.37]	132.1	[2.40]	99.4	[4.99]	-0.094	[-0.68]	-0.133	[-7.83]	0.072	0.590	0.784

The cumulative equity returns Δr_{tA}^j in stock j (denominated in dollars and expressed in percentage points) for different event windows (WS = window size) are regressed on a constant, the change in the local risk premium $[\Sigma^L(w^n - w^o)]_j$, the 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 $[\Sigma^L \Sigma^{Int}(w^n - w^o)]_j$. Formally,

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

The covariance matrix Σ^G is estimated for two 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. Also, $\Sigma^{Int} = \Sigma^G - \Sigma^L$ and $\Sigma^L \Sigma^L = \Sigma^G \Sigma^G - \Sigma^L \Sigma^L$. The event window size is chosen in turn to start WS = 5, 10, 15, 20 trading days prior to December 1, 2000. Panel A reports the coefficients for all stocks, and Panel B only for non-U.S. stocks. Robust and country-clustered adjusted t -values are reported in brackets. The last two columns report the significance level at which equality of the respective coefficients can be rejected.

Table 6
Market integration test for the liquidation of hedging positions

WS	c	[t]	β_2^L	[t]	β_2^{Int}	[t]	R^2	F-Test $\beta_2^L = \beta_2^{Int}$
<i>Panel A: Hedge Liquidation Event (All Stocks, N = 2,291)</i>								
3	0.70	[1.42]	0.019	[4.78]	0.032	[4.40]	0.076	0.007
5	1.89	[2.87]	0.049	[9.98]	0.044	[3.57]	0.139	0.663
7	1.43	[2.24]	0.029	[6.10]	0.042	[3.26]	0.073	0.223
<i>Panel B: Hedge Liquidation Event (Only Added and Deleted Stocks, N = 661)</i>								
3	0.74	[1.77]	0.023	[8.92]	0.022	[2.71]	0.067	0.818
5	2.02	[3.42]	0.058	[14.97]	0.032	[2.10]	0.189	0.053
7	1.52	[2.21]	0.033	[8.28]	0.034	[2.58]	0.075	0.974
<i>Panel C: Hedge Liquidation Event (Non-U.S. Stocks, N = 1,877)</i>								
3	1.50	[4.57]	0.221	[3.32]	0.027	[3.80]	0.111	0.007
5	2.41	[4.90]	0.061	[0.62]	0.043	[3.38]	0.093	0.866
7	1.94	[3.25]	0.109	[1.07]	0.039	[2.78]	0.071	0.527

The cumulative equity returns $\Delta r_{[t_A, t_L]}^j$ in stock j (denominated in dollars and expressed in percentage points) for different event windows (WS = window size) 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,

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

The covariance matrix Σ^G is estimated for two 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. Also, $\Sigma^{Int} = \Sigma^G - \Sigma^L$ and $\Sigma \Sigma^{Int} = \Sigma^G \Sigma^G - \Sigma^L \Sigma^L$. The event window size is chosen in turn to extend over WS = 3, 5, 7 trading days starting on December 4, 2000. Panel A reports the coefficients for all stocks, Panel B only for added and deleted stocks, and Panel C only for non-U.S. stocks. Robust and country-clustered adjusted t -values are reported in brackets. The last column reports the significance level at which equality of the respective coefficients can be rejected.

Table 6 reports the corresponding regression on local versus global pricing for the hedge liquidation period. In the full sample, the coefficient β_2^{Int} is again highly significant and 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 five- and seven-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 that 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 for all three event windows. This is not surprising, given that local marginal arbitrage risk features hardly any cross-sectional variation among non-U.S. stocks. Overall, the hedge liquidation event provides additional support in favor of market integration.

4.4 Integration across emerging and developed markets

The previous section decomposed global beta changes into local (i.e., national) beta changes and complementary international beta changes, and found no

evidence for equity market segmentation along country lines. Alternatively, the global equity market might be segmented between developed and emerging market stocks (marked in Table 1, column (7)). Such market segmentation on a more aggregate level could explain why the coefficients α^{Int} and β^{Int} have explanatory power for the event returns in Tables 5 and 6. This motivates a narrow hypothesis about market segmentation. As before, I test for such segmentation by decomposing the global covariance matrix of all assets Σ^G into a matrix (Σ^H) composed of covariance elements only between developed market and between emerging market stock and a complementary matrix (Σ^{CH}) capturing all (cross-hemisphere) covariances between developed and emerging market stocks. Formally, I define

$$(\Sigma^H)_{ij} = \begin{cases} 0 & \text{if } i, j \text{ combine emerging and developed market stocks,} \\ (\Sigma^G)_{ij} & \text{otherwise.} \end{cases} \quad (23)$$

The components for the premium change and marginal arbitrage risk capturing integration across developed and emerging markets follow as

$$[\Sigma^{CH}(w^n - w^o)]_j = [\Sigma^G(w^n - w^o)]_j - [\Sigma^H(w^n - w^o)]_j, \quad (24)$$

$$[\Sigma \Sigma^{CH}(w^n - w^o)]_j = [\Sigma^G \Sigma^G(w^n - w^o)]_j - [\Sigma^H \Sigma^H(w^n - w^o)]_j, \quad (25)$$

respectively. A test of market integration between emerging and developed equity markets then consists of the cross-sectional regression

$$\begin{aligned} \Delta r_1^j = & c + \alpha_1^H [\Sigma^H(w^n - w^o)]_j + \alpha_1^{CH} [\Sigma^{CH}(w^n - w^o)]_j \\ & + \beta_1^H [\Sigma^H \Sigma^H(w^n - w^o)]_j + \beta_1^{CH} [\Sigma \Sigma^{CH}(w^n - w^o)]_j + \mu_j, \end{aligned} \quad (26)$$

where $\alpha_1^H = \alpha_1^{CH}$ and $\beta_1^H = \beta_1^{CH}$ correspond to the null hypothesis of market integration across developed and emerging markets and $\alpha_1^{CH} = 0$ and $\beta_1^{CH} = 0$ correspond to the null hypothesis of complete market segmentation.

The corresponding evidence for the speculative position buildup is presented in Table 7. The coefficient estimates here are quantitatively similar to those in Table 5. An economic interpretation suggests that the respective return effects are again large. A coefficient estimate of $\alpha_1^H = 147.1$ for the 20-day window implies, *ceteris paribus*, a stock return effect of 7.5% ($= 147.1 \times 0.051$) for a change in the beta of 1 SD. According to Panel E of Table 2, the SD for the arbitrage risk term $[\Sigma^H \Sigma^H(w^n - w^o)]_j$ is given by 60.53. The coefficient estimate of $\beta_1^H = -0.188$ then implies that a 1 SD arbitrage risk change modifies the stock price by 11.38% over the 20-day event window. The coefficients α_1^H and α_1^{CH} are both positive and of similar magnitude;

Table 7
Market integration across emerging and developed markets

WS	c	$[t]$	α_1^H	$[t]$	α_1^{CH}	$[t]$	β_1^H	$[t]$	β_1^{CH}	$[t]$	R^2	F-Test $\alpha_1^H = \alpha_1^{CH}$	F-Test $\beta_1^H = \beta_1^{CH}$
<i>Panel A: Position Buildup Event (All Stocks, N = 2,291)</i>													
5	1.37	[2.74]	19.7	[1.25]	78.6	[3.20]	-0.041	[-2.55]	-0.096	[-5.46]	0.104	0.068	0.039
10	-0.24	[-0.31]	81.6	[3.19]	79.4	[2.20]	-0.100	[-4.88]	-0.097	[-3.62]	0.089	0.962	0.938
15	-2.05	[-1.56]	98.9	[2.78]	108.2	[3.01]	-0.103	[-3.27]	-0.108	[-3.47]	0.051	0.840	0.892
20	-1.74	[-1.18]	147.1	[3.75]	98.0	[1.97]	-0.188	[-5.18]	-0.123	[-2.90]	0.125	0.362	0.180
<i>Panel B: Position Buildup Event (Non-U.S. Stocks, N = 1,877)</i>													
5	1.22	[2.24]	16.6	[0.79]	76.4	[3.11]	-0.025	[-1.61]	-0.092	[-4.86]	0.086	0.102	0.014
10	-0.63	[-0.84]	73.6	[1.98]	72.2	[1.97]	-0.089	[-3.53]	-0.091	[-3.30]	0.063	0.980	0.963
15	-3.00	[-2.34]	68.9	[1.52]	98.5	[2.79]	-0.079	[-2.22]	-0.100	[-3.32]	0.036	0.598	0.663
20	-2.95	[-2.35]	111.4	[2.39]	86.1	[1.74]	-0.156	[-4.14]	-0.113	[-2.35]	0.075	0.685	0.404
<i>Panel C: Position Buildup Event (Emerging Market Stocks, N = 771)</i>													
5	1.46	[3.89]	88.6	[1.12]	15.09	[1.03]	0.060	[0.65]	-0.023	[-1.80]	0.131	0.381	0.389
10	-0.01	[-0.01]	116.1	[1.05]	34.50	[1.67]	-0.009	[-0.07]	-0.051	[-2.87]	0.103	0.493	0.755
15	-4.23	[-6.26]	224.2	[1.42]	65.5	[2.60]	-0.230	[-1.23]	-0.065	[-3.07]	0.036	0.347	0.399
20	-4.02	[-5.09]	206.5	[1.20]	49.4	[1.48]	-0.218	[-1.06]	-0.075	[-2.72]	0.052	0.394	0.505

The cumulative equity returns Δr_{tA}^j in stock j (denominated in dollars and expressed in percentage points) for different event windows (WS = window size) are regressed on a constant, the change in the risk premium $[\Sigma^H(w^n - w^o)]_j$ due to weight changes within the stock's hemisphere of either emerging or developed market stocks, the incremental effect to the global risk premium change $[\Sigma^{CH}(w^n - w^o)]_j$, the marginal arbitrage risks $[\Sigma^H \Sigma^H(w^n - w^o)]_j$ for the arbitrage portfolios specific to developed and emerging markets, and the incremental international arbitrage risk to the global arbitrage risk, $[\Sigma \Sigma^{CH}(w^n - w^o)]_j$. Formally,

$$\Delta r_{tA}^j = c + \alpha_1^H [\Sigma^H(w^n - w^o)]_j + \alpha_1^{CH} [\Sigma^{CH}(w^n - w^o)]_j + \beta_1^H [\Sigma^H \Sigma^H(w^n - w^o)]_j + \beta_1^{CH} [\Sigma \Sigma^{CH}(w^n - w^o)]_j + \mu_j.$$

The global covariance matrix Σ^G is estimated for two years of weekly dollar stock returns for the period of July 1, 1998, to July 1, 2000. The matrix Σ^H is obtained by setting to zero all covariances for stock pairs with one emerging and one developed market stock. Also, $\Sigma^{CH} = \Sigma^G - \Sigma^H$ and $\Sigma \Sigma^{CH} = \Sigma^G \Sigma^G - \Sigma^H \Sigma^H$. The event window size is chosen in turn 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 non-U.S. stocks, and Panel C only for emerging market stocks. Robust and country-clustered adjusted t -values are reported in brackets. The last two columns report the significance level at which equality of the respective coefficients can be rejected.

the F -test cannot reject the hypothesis that both coefficients are the same for each of the four event windows and each of the two samples. Index weight changes by developed market stocks produce (via global beta changes) similar return effects on both emerging and developed market equity. The price effect between developed and emerging market stocks is also confirmed for risk hedging. The coefficients β_1^H and β_1^{CH} are both negative and of very similar size, which supports the market integration hypothesis. In contrast, the market segmentation hypothesis with $\beta_1^{CH} = 0$ is clearly rejected by the data. This rejection also applies to the subsample of non-U.S. stocks (Table 7, Panel B) and for the still smaller sample of only emerging market stocks (Table 7, Panel C). This is direct evidence against full market segmentation between developed and emerging market stocks in the MSCI stock universe. However, it does not preclude the possibility that a subset of emerging market stocks could be segmented from the global equity market. This issue is explored in detail in the next section.

4.5 Market integration by cross-listing and liquidity

Some emerging market stocks might be more integrated into the market for global risk trading than others. In particular, cross-listing of emerging market stocks might facilitate global equity risk sharing.²⁰ Therefore, the index weight change of developed market stocks might exercise a more pronounced effect on an emerging market stock if it is cross-listed than otherwise. To test this hypothesis, the emerging market stocks are sorted into a cross-listed group (*List+*)—with secondary listings in either the NYSE, Nasdaq, or London Stock Exchange—and a complementary set (*List-*) without such a cross-listing. The sample contains 127 (16.5%) cross-listed stocks among the 771 emerging market stocks. Their median index weight w^o is roughly twice that of stocks without cross-listing.

A second potential dimension of market segmentation is illiquidity. Conventional liquidity measures are difficult to calculate for emerging market stocks because of a lack of data on stock turnover. Lesmond (2005) and Bekaert, Harvey, and Lundblad (2007) therefore suggest the percentage of daily zero returns as an illiquidity proxy. The two-year period from July 1, 1998, to July 1, 2000, is used to calculate the percentage of zero returns, ZR_j , and a stock liquidity proxy is defined as $Liq^j = \ln(1 - ZR_j)$. Illiquid stocks tend to be emerging market stocks. Approximately 50% of all emerging market stocks exhibit more than 14.5% zero return days, compared with less than 9% for developed market stocks. Any liquidity-based market segmentation should therefore concern mostly emerging market stocks. To test for liquidity-based market segmentation, the emerging market stocks are sorted into the 50% least

²⁰ Similarly, we may test if cross-listing of developed market stocks is increasing market integration as well. I focus here on emerging market stocks because their cross-listing effect is plausibly more pronounced.

liquid stocks ($Liq-$) and the 50% most liquid stocks ($Liq+$). The least liquid stocks tend to be smaller, and their median index weight w^o is only 41% of the median weight of the most liquid stocks. Cross-listing has a weak positive correlation with liquidity: Of the 127 cross-listed stocks, 82 (64%) are in the more liquid subsample.

The next step is to further decompose the covariance matrix Σ^{CH} in Equation (24). This matrix captures market integration between developed and emerging market stocks and consists of (nonzero) covariance elements only for pairs of developed and emerging market stocks. Both a listing-based and a liquidity-based decomposition of Σ^{CH} are considered. As before, I define a covariance matrix Σ^{H2} with zero elements for the stock pairs characterized by market segmentation. Formally,

$$(\Sigma^{H2})_{ij} = \begin{cases} 0 & \text{if } i, j \text{ combine developed and} \\ & \text{non-cross-listed emerging market stocks,} \\ (\Sigma^G)_{ij} & \text{otherwise.} \end{cases} \quad (27)$$

Compared with Σ^H in Section 4.4, the matrix Σ^{H2} contains fewer zero elements because the segmentation assumption is confined to emerging market stocks without cross-listing. Let $\Sigma^{List-} = \Sigma^G - \Sigma^{H2}$ denote the difference from the full global covariance matrix Σ^G . The elements of Σ^{List-} are equal to Σ^{CH} *except* for those elements involving a cross-listed emerging market stock. The complementary component of Σ^{CH} is then defined by $\Sigma^{List+} = \Sigma^{CH} - \Sigma^{List-}$ and captures the interaction between developed and cross-listed emerging market stocks.

An analogous liquidity-based decomposition of Σ^{CH} follows as

$$(\Sigma^{H3})_{ij} = \begin{cases} 0 & \text{if } i, j \text{ combine developed and illiquid} \\ & \text{emerging market stocks,} \\ (\Sigma^G)_{ij} & \text{otherwise,} \end{cases} \quad (28)$$

with $\Sigma^{Liq-} = \Sigma^G - \Sigma^{H3}$ and $\Sigma^{Liq+} = \Sigma^{CH} - \Sigma^{Liq-}$. A similar decomposition is applied to the squared matrix $\Sigma \Sigma^{CH}$. Formally,

$$\Sigma \Sigma^{List-} = \Sigma^G \Sigma^G - \Sigma^{H2} \Sigma^{H2}, \quad \Sigma \Sigma^{List+} = \Sigma \Sigma^{CH} - \Sigma^{List-} \Sigma^{List-}; \quad (29)$$

$$\Sigma \Sigma^{Liq-} = \Sigma^G \Sigma^G - \Sigma^{H3} \Sigma^{H3}, \quad \Sigma \Sigma^{Liq+} = \Sigma \Sigma^{CH} - \Sigma^{Liq-} \Sigma^{Liq-}. \quad (30)$$

This implies the following decomposition for the terms capturing market integration between developed and emerging markets:

$$[\Sigma^{CH}(w^n - w^o)]_j = [\Sigma^{List+}(w^n - w^o)]_j + [\Sigma^{List-}(w^n - w^o)]_j, \quad (31)$$

$$[\Sigma^{CH}(w^n - w^o)]_j = [\Sigma^{Liq+}(w^n - w^o)]_j + [\Sigma^{Liq-}(w^n - w^o)]_j; \quad (32)$$

$$\begin{aligned} [\Sigma \Sigma^{CH}(w^n - w^o)]_j &= [\Sigma \Sigma^{List+}(w^n - w^o)]_j \\ &\quad + [\Sigma^{List-} \Sigma^{List-}(w^n - w^o)]_j, \end{aligned} \quad (33)$$

$$\begin{aligned} [\Sigma \Sigma^{CH}(w^n - w^o)]_j &= [\Sigma \Sigma^{Liq+}(w^n - w^o)]_j \\ &\quad + [\Sigma^{Liq-} \Sigma^{Liq-}(w^n - w^o)]_j. \end{aligned} \quad (34)$$

Substitution into Equation (26) provides two new regression specifications that separate the effect of the developed market weight changes on emerging market stocks by cross-listing and by liquidity. Equality of coefficients for the decomposed terms on the right-hand side of Equation (26) implies equal market integration for emerging market stocks across listing and liquidity characteristics. The relevant regression coefficients are reported in Table 8, where Panel A provides results on the role of cross-listing and Panel B on the role of liquidity. The next-to-last column provides an F -test for the hypothesis that cross-listed and not cross-listed (or liquid and illiquid) stocks both show event returns in accordance with their global beta change. The F -test in the last column tests equality of the return impact for global arbitrage risk.

The coefficient α_1^{List+} in Panel A of Table 8 characterizes the pricing effect related to market integration between developed and cross-listed emerging market stocks. The estimate $\alpha_1^{List+} = 432.4$ in Panel A for the 20-day window has a t -value of 4.13. The implied return effect is 6.92% ($= 432.4 \times 0.016$) for a beta change of 1 SD. This economic magnitude is relatively close to the 9.35% and 7.5% return effect found for α_1^L and α_1^H in Tables 5 and 7, respectively. In contrast, the corresponding estimate α_1^{List-} for non-cross-listed emerging market stocks is small and statistically insignificant. This means that weight changes of developed market stocks have a large effect on the event returns of cross-listed stocks but not on emerging market stocks without a cross-listing. The F -test rejects equality of the coefficients at the 1% level for all four event windows, which indicates that the international market integration of emerging market stocks differs according to their listing characteristics. The second comparison concerns the coefficients β_1^{List+} and β_1^{List-} for the price impact of the hedging demand. The coefficient β_1^{List+} is significantly larger than β_1^{List-} . The hedging effect is therefore stronger for cross-listed stocks than for those without a secondary listing, as confirmed by the F -test in the last column.

Table 8
Emerging market stock integration by cross-listing and liquidity characteristics

WS	α_{List+}	α_{List-}	β_{List+}	β_{List-}	R^2	F-Test $\alpha_{List+} = \alpha_{List-}$	F-Test $\beta_{List+} = \beta_{List-}$
<i>Panel A: Position Buildup Event with Cross Listing Decomposition (All Stocks, N=2,291)</i>							
5	219.0	28.8	-0.210	-0.036	0.120	0.001	0.001
10	317.7	-5.2	-0.293	0.004	0.114	0.000	0.000
15	337.8	27.0	-0.296	-0.012	0.067	0.004	0.003
20	432.4	-20.3	-0.396	0.017	0.148	0.002	0.002
<i>Panel B: Position Buildup Event with Liquidity Decomposition (All Stocks, N=2,291)</i>							
WS	α_{Liq+}	α_{Liq-}	β_{Liq+}	β_{Liq-}	R^2	F-Test $\alpha_{Liq+} = \alpha_{Liq-}$	F-Test $\beta_{Liq+} = \beta_{Liq-}$
5	103.8	12.8	-0.127	-0.003	0.116	0.006	0.001
10	127.8	-25.1	-0.147	0.036	0.101	0.016	0.004
15	157.5	13.7	-0.153	0.002	0.057	0.073	0.060
20	172.9	-56.5	-0.197	0.067	0.137	0.014	0.007

The regressions in Table 7 are repeated under two different decompositions of the terms $[\Sigma^{CH}(w^n - w^o)]_j$ and $[\Sigma^{Liq}(w^n - w^o)]_j$ capturing market integration between developed and emerging market stocks. Panel A reports a cross-listing-based decomposition, and Panel B a liquidity-based decomposition. The former defines a covariance Σ^{Liq+} with elements equal to Σ^{CH} if the emerging market stock is cross-listed and zero otherwise; the complementary matrix follows as $\Sigma^{Liq-} = \Sigma^{CH} - \Sigma^{Liq+}$ and contains covariance elements for which the emerging market stocks are not cross-listed. The liquidity-based decomposition defines a covariance Σ^{Liq+} with elements equal to Σ^{CH} if the emerging market stock is among the 50% most liquid emerging market stocks and zero otherwise; the complementary matrix here is $\Sigma^{Liq-} = \Sigma^{CH} - \Sigma^{Liq+}$. An analogous decomposition is applied to the squared covariance matrix Σ^{Liq+} . The reported coefficients $\alpha_{List+}, \alpha_{List-}, \beta_{List+}, \beta_{List-}$ in Panel A correspond to the regressors $[\Sigma^{Liq+}(w^n - w^o)]_j, [\Sigma^{Liq-}(w^n - w^o)]_j, [\Sigma^{Liq+}(w^n - w^o)]_j, [\Sigma^{Liq-}(w^n - w^o)]_j$, respectively, and the coefficients $\alpha_{Liq+}, \alpha_{Liq-}, \beta_{Liq+}, \beta_{Liq-}$ in Panel B correspond to the regressors $[\Sigma^{Liq+}(w^n - w^o)]_j, [\Sigma^{Liq-}(w^n - w^o)]_j, [\Sigma^{Liq+}(w^n - w^o)]_j$ and $[\Sigma^{Liq-}(w^n - w^o)]_j$, respectively. Financial market integration of emerging market stocks without cross-listing (or low liquidity) implies $\alpha_{List+} = \alpha_{List-} = \alpha_{Liq+} = \alpha_{Liq-} = \beta_{List+} = \beta_{List-} = \beta_{Liq+} = \beta_{Liq-} = 0$. Coefficients for the constant term and the additional regressors $[\Sigma^H(w^n - w^o)]_j$ and $[\Sigma^H(w^n - w^o)]_j$ are not reported. All covariance matrices are estimated for two years of weekly dollar stock returns for the period of July 1, 1998, to July 1, 2000. The event window size is chosen in turn to start WS = 5, 10, 15, 20 trading days prior to December 1, 2000. Robust and country-clustered adjusted *t*-values are reported in brackets. The last two columns report the significance level at which equality of the respective coefficients can be rejected.

Panel B of Table 8 compares the pricing effects of market integration of liquid and illiquid emerging market stocks, where α_1^{Liq+} (α_1^{Liq-}) averages the return effects related to liquid (illiquid) emerging market stocks. The estimate $\alpha_1^{Liq+} = 172.9$ for the 20-day window corresponds to a 4.15% ($= 172.9 \times 0.024$) return effect for a 1 SD beta change. The corresponding coefficient α_1^{Liq-} for illiquid emerging market stocks is statistically insignificant. Generally, benchmark weight changes of developed market stocks induce a smaller return effect for the 50% most illiquid emerging market stocks, since α_1^{Liq-} is lower than α_1^{Liq+} . The difference is (in all but one case) statistically significant at the 3% level. The coefficient α_1^{Liq-} is statistically insignificant, so the market segmentation hypothesis cannot be rejected for illiquid stocks. However, this hypothesis is strongly rejected for the 50% most liquid stocks. The larger return effect for the more liquid stocks also contrasts with the traditional “price pressure” interpretation based on individual stock weight changes. Under the (univariate) price pressure hypothesis, index weight changes should generate larger returns for more illiquid stocks—but not the opposite. Overall, the evidence shows that relatively liquid emerging market stocks are integrated into the global equity market. In contrast, the 50% most illiquid emerging market stocks appear to be segmented in terms of their risk pricing. This latter finding applies to 385 of the 2,291 stocks in the MSCI sample.

5. Robustness Issues

A variety of robustness issues can be raised with respect to the evidence presented in previous sections. Here, I discuss (i) concerns about potential regressor collinearity; (ii) the role of direct “price pressure” and liquidity effects as alternative explanations for the cross-sectional return patterns; and (iii) inference issues with respect to the covariance matrix Σ .

Regressor collinearity poses an inference problem under high correlation of the independent variables and a small number of observations. The two main regressors, $[\Sigma^G(w^n - w^o)]_j$ and $[\Sigma^G \Sigma^G(w^n - w^o)]_j$, have a correlation of 0.931 for the entire sample of 2,291 stocks. The correlation is slightly higher at 0.943 for the subsample of 661 added and deleted stocks and drops to 0.878 for the 1,877 non-U.S. stocks. But these correlations should not pose a collinearity problem in light of the large number of observations. Also, the coefficients α_1 and β_1 are of similar magnitude in the full sample and in the subsample of non-U.S. stocks for which the regressor correlation is lowest.

The equity return effects considered in this article are based on the identification of anticipated beta changes and risk-hedging concerns. This contrasts with much of the literature on index inclusions and exclusions that considers direct (univariate) price pressure effects. Such price pressure may simply be proportional to the percentage weight change; that is, $PP^j = 2(w^n - w^o)_j / (w^n + w^o)_j$. I therefore augment the regression in Table 3 by

this price pressure proxy. The finance literature has highlighted the role of stock liquidity for (short-term) stock price behavior, which suggests liquidity measures as additional regression controls. Following [Bekaert, Harvey, and Lundblad \(2007\)](#), I use again the liquidity proxy $Liq^j = \ln(1 - ZR^j)$, where ZR^j measures the percentage of days with zero stock return. [Table 9](#) presents the augmented regression results. The specification in Panel C allows also for an interaction between price pressure and liquidity effects. Although the short five-day event window shows modest liquidity effects, the overall regression results are unchanged. In particular, price pressure as measured by PP^j cannot account for the cross-sectional price variation. A similar result (not reported) is obtained if I exclude from the regression all added and deleted stocks, for which the price pressure proxy PP^j degenerates to 2 and -2 , respectively. The interaction term $PP^j \times Liq^j$ should capture liquidity improvement if percentage weight changes improve liquidity especially for illiquid stocks. But the interaction term is statistically insignificant. Neither direct price pressure nor liquidity effects (nor their interaction) account for the cross-sectional return pattern observed prior to the announcement of the index change. On the other hand, the return-seeking component and the risk-hedging component both retain their economic and statistical significance level under these controls.

Estimation error with respect to the covariance matrix Σ is another potential concern. This problem tends to be particularly severe if the covariance matrix is inverted. The empirical inference in this article does not rely on such a matrix inversion. Rather, the row elements of the matrix Σ are averaged when multiplied by the vector of index weight changes $w^n - w^o$. This implies that estimation errors with respect to each matrix element are also averaged. Even though I am calculating a high-dimensional ($2,291 \times 2,291$) matrix, the averaging implies that effectively only a vector with 2,291 row elements is estimated. A same logic applies also to the marginal arbitrage risk $[\Sigma \Sigma (w^n - w^o)]_j$, where post-multiplication by $w^n - w^o$ again overcomes the curse of high dimensionality with respect to $\Sigma \Sigma$.

Another robustness test consists of estimating a factor model for the covariance matrix. A fitted covariance matrix is estimated based on the first 20, 40, or 60 principal components of the covariance matrix. The corresponding factor models capture (respectively) 50.6%, 74.8%, or 92.1% of the total two-year return variation. The regressions and market integration tests in [Table 5](#) are repeated under these fitted covariance estimates (and are available on the author's website; the regression outcomes are similar under either specification of the covariance matrix). As for the original sample covariance, the coefficients α_1^{Int} and β_1^{Int} are highly significant and F -tests fail to reject the market integration hypothesis. Reproducing the regression results under a lower-dimensional factor representation of the covariance Σ is further evidence of robustness. Only when the number of factors decreases below 20 does the model fit deteriorate.

Table 9
Price pressure and liquidity effects for the speculative position buildup

WS	c	$[t]$	γ_1	$[t]$	γ_2	$[t]$	γ_3	$[t]$	α_1	$[t]$	β_1	$[t]$	R^2
<i>Panel A: Position Buildup Event with Price Pressure Controls (All Stocks, N = 2,291)</i>													
5	1.56	[3.32]	0.13	[0.71]					41.2	[3.37]	-0.064	[-6.53]	0.095
10	-0.22	[-0.29]	0.18	[0.53]					79.7	[3.99]	-0.099	[-6.39]	0.089
15	-2.02	[-1.56]	0.16	[0.32]					100.9	[3.76]	-0.105	[-4.45]	0.052
20	-1.95	[-1.31]	0.33	[0.51]					122.9	[3.46]	-0.161	[-4.55]	0.119
<i>Panel B: Position Buildup Event with Price Pressure and Liquidity Controls (All Stocks, N = 2,291)</i>													
5	2.44	[4.61]	0.03	[0.14]	6.24	[4.18]			41.7	[3.56]	-0.067	[-7.30]	0.107
10	0.33	[0.34]	0.12	[0.33]	3.94	[1.50]			80.0	[4.07]	-0.100	[-6.43]	0.091
15	-1.10	[-0.69]	0.04	[0.09]	6.49	[1.64]			101.4	[3.90]	-0.108	[-4.51]	0.053
20	-1.65	[-0.86]	0.29	[0.57]	2.10	[0.46]			123.0	[3.47]	-0.162	[-4.45]	0.120
<i>Panel C: Position Buildup Event with All Controls (All Stocks, N = 2,291)</i>													
5	2.47	[4.82]	0.08	[0.29]	6.54	[3.77]	0.40	[0.36]	41.7	[3.56]	-0.067	[-7.35]	0.107
10	0.51	[0.53]	0.44	[0.94]	5.56	[2.14]	2.22	[1.67]	80.0	[4.09]	-0.101	[-6.48]	0.093
15	-0.85	[-1.58]	0.50	[0.75]	8.81	[2.09]	3.17	[1.65]	101.4	[3.92]	-0.109	[-4.53]	0.058
20	-1.49	[-0.79]	0.59	[0.86]	3.60	[0.77]	2.05	[1.21]	123.0	[3.48]	-0.163	[-4.46]	0.120

The regression in Table 3 is repeated using additional control variables. A price pressure proxy is defined by the stock's percentage weight change $PP^j = 2(w^n - w^o)_j / (w^n + w^o)_j$ and a liquidity proxy for each stock is given by $Liq^j = \ln(1 - ZR^j)$, where ZR^j denotes the percentage of zero return days over a prior two-year period. We also allow for an interaction term between both variables. The cumulative event returns Δr_{tA}^j (denominated in dollars and expressed in percentage points) over different even windows (WS = window size) is regressed on a constant, the change in the risk premium $[\Sigma^G(w^n - w^o)]_j$, and the arbitrage risk $[\Sigma^G(w^n - w^o)]_j$ of each stock j . Formally,

$$\Delta r_{tA}^j = c + \gamma_1 PP^j + \gamma_2 Liq^j + \gamma_3 (PP^j \times Liq^j) + \alpha_1 [\Sigma^G(w^n - w^o)]_j + \beta_1 [\Sigma^G(w^n - w^o)]_j + \mu_j.$$

The covariance matrix Σ^G is estimated for two years of weekly dollar stock returns for the period of July 1, 1998, to July 1, 2000. The event window size is chosen in turn to start WS = 5, 10, 15, 20 trading days prior to December 1, 2000. Panel A reports the coefficients for price pressure controls, Panel B for stock liquidity controls, and Panel C for both controls as well as their interaction term. Robust and country-clustered adjusted t -values are reported in brackets.

6. Conclusion

This article argues that large-scale index revisions can modify the market portfolio. Such a change in the market portfolio (or market benchmark) has highly predictable and testable implications for asset pricing. Since the beta of a stock is proportional to the covariance of the stock return with the market benchmark, the latter's change should also change the stock beta. For constant cash flow expectations, beta changes generate cross-sectional event returns around the announcement of the index change. *Ceteris paribus*, stocks with a beta increase (decrease) should experience a negative (positive) excess return. Such an event-based asset pricing test amounts to testing asset pricing models in market benchmark differences. Hence, this test does not require explicit specification of the market benchmark itself—provided the benchmark change is clearly defined by the index change. This avoids the joint hypothesis problem of testing both a benchmark selection and a pricing model that characterizes much of the previous asset pricing literature.

The index revision considered in this article involves the substantial redefinition of the MSCI global equity index announced in December 2000. The unprecedented scope of the index revision yields a sample of 2,291 stocks for which beta changes can be calculated. Moreover, the global nature of the benchmark change allows for discrimination between global and local beta changes. This suggests a simple test of global versus local asset pricing. If the global beta changes are sufficient to characterize excess returns around the announcement of the index revision, then we face globally integrated risk pricing, in which case a global market benchmark is appropriate. Alternatively, market segmentation into local markets implies that local beta changes are a better explanatory variable for event returns.

However, implementation of this simple empirical strategy requires additional considerations. Risk arbitrageurs might anticipate an index revision and thereby trigger intense speculative front-running of the index announcement. Such speculative trading may generate additional confounding effects for event returns. A simple model of speculative arbitrage is used to understand how this occurs. The optimal trading strategy for mean–variance speculators consists of a trade-off between higher expected returns and lower arbitrage risk. The optimal arbitrage portfolio can be represented as a linear combination of a return-seeking portfolio and a risk-hedging portfolio. The hedging demand generates an additional correlated event return pattern. For this reason, the econometric strategy must control for the price impact of transitory speculative hedging in order to identify the return effect of beta changes.

Two principal empirical findings can be highlighted. First, a speculative model that accounts for hedging demands provides a much better fit to the observed MSCI event returns than do alternative models. In the run-up to the announcement, both the return-seeking component and the risk-hedging component are highly significant determinants of the cross-section of returns,

and also have the predicted signs. The subsequent period features a positive return effect for the risk-hedging component, as entailed by the liquidation of such hedging positions. These findings are robust to variations of the event window size and also extend to various subsamples. In contrast, alternative explanations like direct price pressure or liquidity effects cannot account for the cross-section of event returns.

Second, the MSCI event study provides evidence of globally integrated risk pricing. A country-based market segmentation hypothesis can be rejected because event returns are best captured by global, not local, beta changes. Therefore, asset pricing models that use a global benchmark are more appropriate than models based on a local market benchmark. A similar conclusion is reached with respect to arbitrage risk. The international component of the marginal arbitrage risk is a highly significant pricing factor, which suggests that arbitrage strategies for the MSCI revision were implemented globally. The data reject an even more narrowly framed hypothesis of market segmentation between developed and emerging markets. The pricing effects here support global asset pricing especially for the most liquid and/or cross-listed emerging market stocks. Only those emerging market stocks without a cross-listing or with below-median liquidity show evidence of market segmentation in terms of risk pricing. Overall, the evidence suggests that a global market benchmark is appropriate for the large majority of MSCI stocks.

Appendix

Proof of Proposition 1. Let expectations that incorporate knowledge about the demand shock u be denoted by $\tilde{\mathcal{E}}_t(\cdot)$ and those that do not by $\bar{\mathcal{E}}_t(\cdot)$. The market-clearing conditions then follow as

$$\begin{aligned}
 \lambda(\rho \Sigma \Delta t)^{-1} \bar{\mathcal{E}}_t(p_{t+\Delta t} - p_t) + (1 - \lambda)\gamma \bar{\mathcal{E}}_t(p_{t+\Delta t} - p_t) &= \tilde{S}^o & \text{for } 0 \leq t < t_A \\
 \lambda(\rho \Sigma \Delta t)^{-1} \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t) + (1 - \lambda)\gamma \bar{\mathcal{E}}_t(p_{t+\Delta t} - p_t) &= \tilde{S}^o & \text{for } t_A \leq t < t_L \\
 \lambda(\rho \Sigma \Delta t)^{-1} \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t) + (1 - \lambda)\gamma \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t) &= \tilde{S}^o & \text{for } t_L \leq t < t_u \\
 \lambda(\rho \Sigma \Delta t)^{-1} \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t) + (1 - \lambda)\gamma \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t) &= \tilde{S}^o - u & \text{for } t_u \leq t < T
 \end{aligned} \tag{A1}$$

where the left-hand-side terms in Equation (A1) represent the respective asset demand of the speculators and the liquidity suppliers; \tilde{S} denotes the total asset supply (net of index capital) and $u = \vartheta(w^n - w^o)$ the demand shock of index capital at time t_u . By assumption, arbitrageurs learn about the index change at time $t_A < t_u$, whereas liquidity suppliers do so only at time t_L with $t_A < t_L < t_u$. The expected terminal asset price is identical for both groups and is given by

$$\mathcal{E}_{t=k\Delta t}(p_T) = \mathbf{1} + \sum_{t=\Delta t}^{k\Delta t} \Delta \varepsilon_t. \tag{A2}$$

The expected equilibrium return $r_4 \Delta t = \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t)$ from t to $t + \Delta t$ for $t_u \leq t < T$ follows directly from Equation (A1) as

$$\begin{aligned} r_4 \Delta t &= \tilde{\mathcal{E}}_t(p_{t+\Delta t} - p_t) = [\lambda(\rho \Sigma \Delta t)^{-1} + (1 - \lambda)\gamma I]^{-1} (\tilde{S}^o - u) \\ &= \left[I + (1 - \lambda)\gamma \frac{\rho}{\lambda} \Sigma \Delta t \right]^{-1} \frac{\rho}{\lambda} \Sigma (\tilde{S}^o - u) \Delta t \\ &\approx \left[I - (1 - \lambda)\gamma \frac{\rho}{\lambda} \Sigma \Delta t \right] \frac{\rho}{\lambda} \Sigma (\tilde{S}^o - u) \Delta t \\ &\approx \frac{\rho}{\lambda} \Sigma (\tilde{S}^o - u) \Delta t, \end{aligned} \tag{A3}$$

where I use the approximation $[I + k \Sigma \Delta t]^{-1} \approx I - k \Sigma \Delta t$ for small $k \Delta t$ and ignore terms of higher order. I note that the approximation quality is high ($k \approx 0$) if the weight $1 - \lambda$ of the liquidity suppliers is small.

For the period $t_L \leq t < t_u$, the supply change u is not yet effective; hence, the expected return simplifies to

$$r_3 \Delta t \approx \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t. \tag{A4}$$

The asset price follows (by recursive substitution) as

$$p_t \approx \begin{cases} \mathcal{E}_t(p_T) - (T - t_u)r_4 - (t_u - t)r_3 & \text{for } t_L \leq t < t_u, \\ \mathcal{E}_t(p_T) - (T - t)r_4 & \text{for } t_u \leq t < T. \end{cases} \tag{A5}$$

For the period $t_A \leq t < t_L$, expectations about the correct equilibrium price differ between arbitrageurs, who know about the demand shock u , and liquidity suppliers, who do not. Hence, expectations are given by

$$\tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) = \mathcal{E}_{t_L - \Delta t}(p_T) - (T - t_u)r_4 - (t_u - t_L)r_3, \tag{A6}$$

$$\bar{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) = \mathcal{E}_{t_L - \Delta t}(p_T) - (T - t_u)r_3 - (t_u - t_L)r_3, \tag{A7}$$

and the valuation difference between liquidity suppliers and arbitrageurs follows as

$$\bar{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) - \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) = (T - t_u)(r_4 - r_3) = -\frac{\rho}{\lambda} \Sigma u(T - t_u). \tag{A8}$$

The market-clearing condition in Equation (A1) for $t = t_L - \Delta t$ implies (under substitution of Equation (A8)) that

$$\begin{aligned} p_{t_L - \Delta t} &= [\lambda(\rho \Sigma \Delta t)^{-1} + (1 - \lambda)\gamma I]^{-1} \\ &\quad \times [-\tilde{S}^o + \lambda(\rho \Sigma \Delta t)^{-1} \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) + (1 - \lambda)\gamma \bar{\mathcal{E}}_{t_L - \Delta t}(p_{t_L})] \\ &= \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) - \frac{\rho}{\lambda} \Sigma \Delta t [I + (1 - \lambda)\gamma \frac{\rho}{\lambda} \Sigma \Delta t]^{-1} [\tilde{S}^o - (1 - \lambda)\gamma (T - t_u)(r_4 - r_3)]. \end{aligned} \tag{A9}$$

Using the approximation $[I + k \Sigma \Delta t]^{-1} \approx I - k \Sigma \Delta t$ with $k = (1 - \lambda)\gamma \frac{\rho}{\lambda}$ and ignoring terms of higher order yields

$$\begin{aligned} p_{t_L - \Delta t} &\approx \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) - \frac{\rho}{\lambda} \Sigma \Delta t [I - (1 - \lambda)\gamma \frac{\rho}{\lambda} \Sigma \Delta t] [\tilde{S}^o - (1 - \lambda)\gamma (T - t_u)(r_4 - r_3)] \\ &\approx \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) - \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t - (1 - \lambda)\gamma \left(\frac{\rho}{\lambda} \right)^2 \Sigma \Sigma u(T - t_u) \Delta t. \end{aligned} \tag{A10}$$

The equilibrium return for $t = t_L - \Delta t$ then follows as

$$r_2 \Delta t = \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L}) - p_{t_L - \Delta t} \approx \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t + (1 - \lambda) \gamma \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma u(T - t_u) \Delta t. \quad (\text{A11})$$

Similarly, for $t = t_L - 2\Delta t$, I obtain the expressions

$$\bar{\mathcal{E}}_{t_L - 2\Delta t}(p_{t_L - \Delta t}) - \tilde{\mathcal{E}}_{t_L - 2\Delta t}(p_{t_L - \Delta t}) = (T - t_u)(r_4 - r_3) + (r_2 - r_3) \Delta t \quad (\text{A12})$$

and

$$\begin{aligned} p_{t_L - 2\Delta t} &\approx \tilde{\mathcal{E}}_{t_L - 2\Delta t}(p_{t_L - \Delta t}) - \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t - (1 - \lambda) \gamma \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma u(T - t_u) \Delta t + A_{\Delta t} \\ &\approx \tilde{\mathcal{E}}_{t_L - \Delta t}(p_{t_L - \Delta t}) - \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t - (1 - \lambda) \gamma \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma u(T - t_u) \Delta t, \end{aligned} \quad (\text{A13})$$

where the cubic term $A_{\Delta t} = (1 - \lambda) \gamma \left(\frac{\rho}{\lambda}\right)^3 \Sigma \Sigma \Sigma u(T - t_u)(\Delta t)^2 \approx 0$ is ignored.²¹ Hence, I find (approximately) the same expected return:

$$r_2 \Delta t = \tilde{\mathcal{E}}_{t_L - 2\Delta t}(p_{t_L - \Delta t}) - p_{t_L - 2\Delta t} \approx \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t + (1 - \lambda) \gamma \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma u(T - t_u) \Delta t. \quad (\text{A14})$$

Repeated substitution (while ignoring the hedging terms in $\left(\frac{\rho}{\lambda}\right)^3 \Sigma \Sigma \Sigma u$) implies, for the equilibrium price,

$$p_t \approx \mathcal{E}_t(p_T) - (T - t_u)r_4 - (t_u - t)r_3 - (t_L - t)r_2 \text{ for } t_A \leq t < t_L. \quad (\text{A15})$$

It is instructive to characterize the speculative positions of the arbitrageurs, which can be stated as

$$x_t^A = \tilde{S}^o - x_t^L = \tilde{S}^o - (1 - \lambda) \gamma \bar{\mathcal{E}}_t(p_{t+\Delta t} - p_t). \quad (\text{A16})$$

Substituting the expectations of the liquidity suppliers, given by $\bar{\mathcal{E}}_t(p_{t+\Delta t}) = \mathcal{E}_t(p_T) - (T - t - \Delta t)r_3$, into Equation (A16) and then using Equation (A15) implies that

$$\begin{aligned} x_t^A &\approx \tilde{S}^o - (1 - \lambda) \gamma r_3 \Delta t + (1 - \lambda) \gamma \frac{\rho}{\lambda} \Sigma u(T - t_u) \\ &\quad - (1 - \lambda)^2 \gamma^2 \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma u(T - t_u)(t_L - t). \end{aligned} \quad (\text{A17})$$

Speculative positions are therefore positively proportional to Σu and negatively proportional to $\Sigma \Sigma u$. The latter term represents the arbitrageurs' hedging position, which decreases linearly as the t_L date approaches. Finally, the price process for the initial period follows as

$$p_t \approx \mathcal{E}_t(p_T) - (T - t)r_3 \text{ for } 0 \leq t < t_A. \quad (\text{A18})$$

The entire price path (adjusted for the expected liquidation value $\mathcal{E}_t(p_T)$) is plotted in Figure 1. To obtain Proposition 1, I determine the price reaction when the speculators learn about the demand

²¹ The price effect of hedging generates additional higher-order hedging demands from the arbitrageurs that I do not account for here. See the technical Online Appendix at www.haraldhau.com for a more detailed discussion of the approximation error.

shock $u = \vartheta(w^n - w^o)$ at time $t = t_A$. This price effect may be written as

$$\begin{aligned}
 p_{t_A} - p_{t_A - \Delta t} &\approx \mathcal{E}_{t_A - \Delta t}(p_T) + \Delta \varepsilon_{t_A - \Delta t} - (T - t_u)r_4 - (t_u - t_L)r_3 - (t_L - t_A)r_2 \\
 &\quad - [\mathcal{E}_{t_A - \Delta t}(p_T) - (T - t_A + \Delta t)r_3] \\
 &= \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t + \frac{\rho}{\lambda} \Sigma u(T - t_u) - (1 - \lambda)\gamma \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma u(T - t_u) \\
 &\quad \times (t_L - t_A) + \Delta \varepsilon_{t_A - \Delta t}. \\
 &= \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t + \alpha_1 \Sigma (w^n - w^o) + \beta_1 \Sigma \Sigma (w^n - w^o) + \Delta \varepsilon_{t_A - \Delta t}. \quad (\text{A19})
 \end{aligned}$$

After subtracting the expected return $\frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t$ for the interval Δt , the excess return is given by

$$\begin{aligned}
 \Delta r_{t=t_A} &= p_{t_A} - p_{t_A - \Delta t} - \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t \approx \alpha_1 \Sigma (w^n - w^o) \\
 &\quad + \beta_1 \Sigma \Sigma (w^n - w^o) + \Delta \varepsilon_{t_A - \Delta t} \quad (\text{A20})
 \end{aligned}$$

with

$$\alpha_1 = \frac{\rho}{\lambda} \vartheta(T - t_u) \quad \text{and} \quad \beta_1 = -(1 - \lambda)\gamma \left(\frac{\rho}{\lambda}\right)^2 \vartheta(T - t_u)(t_L - t_A).$$

The term $\alpha_1 \Sigma (w^n - w^o)$ represents the return-seeking component, and the term $\beta_1 \Sigma \Sigma (w^n - w^o)$ represents the risk-hedging component. The latter is proportional to the duration of the arbitrage position given by $t_L - t_A$.

Proof of Proposition 2. Consider the equilibrium price sequence derived in the proof of Proposition 1. For the trading period $t_A \leq t < t_L$, the expected return between t and $t + \Delta t$ is approximated by

$$\mathcal{E}_t p_{t+\Delta t} - p_t \approx \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t + (1 - \lambda)\gamma \left(\frac{\rho}{\lambda}\right)^2 \Sigma \Sigma (T - t_u)u \Delta t. \quad (\text{A21})$$

The expected excess return over the interval $[t_A, t_L]$ then follows as

$$r_{[t_A, t_L]} = \sum_{t \in [t_A, t_L]} p_t - p_{t-\Delta t} - \frac{\rho}{\lambda} \Sigma \tilde{S}^o \Delta t \approx \beta_2 \Sigma \Sigma (w^n - w^o), \quad (\text{A22})$$

where $\beta_2 = -\beta_1$. This completes the proof of Proposition 2.

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