INTERNATIONAL JOURNAL OF FINANCE AND ECONOMICS
Int. J. Fin. Econ. **12:** 273–285 (2007)
Published online 19 January 2007 in Wiley InterScience
(www.interscience.wiley.com) DOI: 10.1002/ijfe.311

LIMITS TO INTERNATIONAL ARBITRAGE: AN EMPIRICAL EVALUATION

HANS DEWACHTER[a,b] and KRISTIEN SMEDTS[c,*,†,‡]
[a] *CES, KULeuven, Belgium*
[b] *RIFM, Erasmus University Rotterdam, The Netherlands*
[c] *AFI, KULeuven, Belgium*

ABSTRACT

This paper studies international financial integration by testing the law of one price across national borders. We use the distance between national discount factors as an integration measure and analyze the level of cross-border mispricing. The empirical analysis shows that pricing differentials are relatively large in economic terms. This lack of international financial integration is subsequently analyzed in the market micro-finance literature. We find that market characteristics explain a considerable part of the variance in our cross-section of pricing differentials. Copyright © 2007 John Wiley & Sons, Ltd.

JEL CODE: F3; G1

KEY WORDS: Arbitrage; financial integration; EMU

1. INTRODUCTION

In recent years, evidence of imperfect financial market integration has been accumulating. Defining integration in terms of (cross-market) mispricing, many authors have documented large and persistent mispricing anomalies. At least two strands of literature can be distinguished. The first type studies whether prices reflect fundamental values and analyzes the role of national risk factors in the pricing of assets. If financial markets are perfectly integrated internationally, none of the purely national factors are priced. This strand of literature is called absolute valuation as one needs some equilibrium model of asset pricing, e.g. CAPM or IAPT type of models, to determine the fundamental value of assets. The main drawback of this approach is that fundamental values are unobservable and hence that the test of market integration is always a joint test of the validity of the pricing model and market efficiency.

A second strand of literature uses no restriction on the pricing model, but tests the relative valuation of assets as defined by the law of one price. The advantage of this methodology is that one only needs to know whether identical assets have identical prices. The difficulty is that one is restricted to the class of assets that can be valued relative to a base asset. A typical test examines closed-end funds (see Hardouvelis *et al*., 1994; Bodurtha *et al*., 1995), dual listed companies, so-called Siamese-twin stocks (see Froot and Dabora, 1999; de Jong *et al*., 2003), or equity carve-outs (see Lamont and Thaler, 2003).

While the above-mentioned approaches can be used to measure international financial integration, we employ a technique proposed by Chen and Knez (1995) that overcomes the drawbacks of each of these

*Correspondence to: K. Smedts, Catholic University of Leuven, AFI, Naamsestraat 69, 3000 Leuven, Belgium.
† E-mail: kristien.smedts@econ.kuleuven.be
‡ This work was written while Kristien Smedts was Aspirant of the FWO-Vlaanderen at CES, Catholic University of Leuven.

approaches. By focusing on relative asset pricing, the problem of the joint test is avoided. Moreover, the methodology of Chen and Knez (1995) assesses the relative pricing differentials of any type of asset or portfolio. This extension of the set of relative pricing assessments is important because it allows us to generate sufficiently many data to study, more generally, the size of possible pricing differentials.

The paper proceeds in two steps. First, we examine the law of one price to assess the degree of financial integration. We apply a method proposed by Chen and Knez (1995), which builds on the Hansen–Jagannathan pricing framework (1997). Taking into account the international law of one price condition, this methodology is extended to measure international financial integration. Since this is a general pricing framework, we can measure cross-market integration with a minimum of additional pricing assumptions.[1] We are not restricted to certain types of assets, and mispricing is assessed without reference to a functional asset pricing model.

Subsequently, we proceed by analyzing factors that may explain the observed lack of international financial integration. We concentrate on channels that have been suggested by the micro-finance theory. Violations of the law of one price may exist because arbitrageurs may face impediments not incorporated into the standard pricing theories. We analyze some of the factors suggested in the theoretical literature such as market value, market volatility and market activity. We find that most of these market characteristics explain (in cross-section) a significant part of the variance of the pricing differentials across markets.

The remainder of the paper is organized as follows. In Section 2, we introduce the Chen and Knez (1995) methodology that measures the degree of mispricing across markets. Building on their methodology we extend this measurement theory to an international financial environment and concentrate on the cross-market mispricing measures. Subsequently, Section 3 presents the empirical analysis on international financial integration between US and European financial markets. These pair-wise measures of segmentation are then used as an input in the second part of Section 3 where we try to explain the degree of market segmentation by means of the above listed market factors. Finally, Section 4 concludes.

2. MEASURING FINANCIAL INTEGRATION

In this section, we extend the integration metric introduced by Chen and Knez (1995) (CK) to an international setting. The CK-measure of market integration is appealing as it is not tied to any specific asset pricing model. The only underlying principle of the measure is the law of one price: identical assets, sold in any market, should sell at the same price. Deviations between two market prices then suggests deviations of the law of one price and the existence of cross-market segmentation. We use this integration metric to measure international financial integration by studying the law of one price in an international context.

2.1. The law of one price in an international market

Consider a set of asset prices on a specific financial market $P_{i,t}$, for $i = 1, \ldots, N$ with (stochastic) future payoffs $X_{i,t+\tau}$. If the law of one price is satisfied within that market, there exists a pricing kernel $m_{t+\tau}$ such that:

$$P_{i,t} = E_t[m_{t+\tau} X_{i,t+\tau}] \tag{1}$$

Analogously, we introduce a second financial market with different numeraire (different currency) by introducing a pricing kernel $m^*_{t+\tau}$ and a set of prices $P^*_{i,t}$ with stochastic payoffs $X^*_{i,t+\tau}$, $i = 1, \ldots, N^*$. Again assuming that the law of one price holds, implies the existence of the pricing kernel $m^*_{t+\tau}$ such that:

$$P^*_{i,t} = E_t[m^*_{t+\tau} X^*_{i,t+\tau}] \tag{2}$$

Equations (1) and (2) imply that the law of one price holds in each financial market separately. Extending the law of one price across markets requires an additional restriction on the exchange rate dynamics. A sufficient condition to ensure international no-arbitrage is the so-called complete market assumption

DOI: 10.1002/ijfe

formulation of the exchange rate dynamics (see Brandt *et al.*, 2004; Backus *et al.*, 2001):

$$\frac{S_{t+\tau}}{S_t} = \frac{m^*_{i,t+\tau}}{m_{i,t+\tau}} \tag{3}$$

where S_t denotes the unit price of foreign currency in terms of domestic currency. This condition is sufficient to extend the law of one price to international investment opportunities. Importantly, the law of one price in the international market is trivially satisfied for the exchange rate as defined in (3). Moreover, if markets are complete, this determination is unique. If markets are incomplete, the definition is not unique, since the exchange rate is not fully characterized by the minimum-variance discount factors. That is, the exchange rate might deviate from (3) as (see Brandt and Santa-Clara, 2002):

$$\frac{S_{t+\tau}}{S_t} = \frac{m^*_{i,t+\tau}}{m_{i,t+\tau}} O_{t+\tau} \tag{4}$$

where $O_{t+\tau}$ is a martingale that is orthogonal to both the home and the foreign stochastic discount factor. Keeping this in mind, failure of the international law of one price (3) does not necessarily correspond to market segmentation, but might be caused by market incompleteness. However, when markets are incomplete, one can always choose the exchange rate as defined in (4) such that the law of one price holds between any two countries, and thus that markets are integrated.

2.2. *Measure for international financial integration*

International financial integrated markets do not allow for cross-border failures of the law of one price. Based on this definition of financial integration, we use the above pricing equations, (1)–(3), to measure the degree of international financial integration. This framework is used to test for the equality of pricing kernels across markets. More specifically, we extend the home market by the set of exchange rate adjusted payoffs from the foreign market and test for equality of pricing characteristics of equivalent payoffs. Denote the domestic payoff space by $\mathscr{X}_{t+\tau}$, with $\sum_{i=1}^{N} \alpha_i X_{i,t+\tau} \in \mathscr{X}_{t+\tau}$ for $\alpha_i \in \mathbb{R}$ and construct the home currency denominated foreign payoff space by $\mathscr{X}^f_{t+\tau}$ with $\sum_{i=1}^{N^*} \alpha_i (X^*_{i,t+\tau} \cdot S_{t+\tau}) \in \mathscr{X}^f_{t+\tau}$. Absence of cross-border segmentation implies that a single pricing kernel prices both the home and foreign sets of payoffs.

To obtain an operational measure for international financial integration, we compute the minimal distance between the sets of pricing kernels $\mathscr{M}_{t+\tau}$ and $\mathscr{M}^f_{t+\tau}$ defined, respectively, on $\mathscr{X}_{t+\tau}$ and $\mathscr{X}^f_{t+\tau}$. Moreover, by normalizing the payoff space to payoffs with unit norm, we obtain a distance measure that can be compared across portfolios. Define the space of gross returns obtainable in the home market by $\mathscr{G}_{t+\tau}$ with $\sum_{i=1}^{N} \alpha_i X_{i,t+\tau} (\sum_{i=1}^{N} \alpha_i P_{i,t})^{-1} \in \mathscr{G}_{t+\tau}$ and the space of home currency denominated gross returns obtainable from foreign investments by $\mathscr{G}^f_{t+\tau}$ with $\sum_{i=1}^{N^*} \alpha_i S_{t+\tau} X^*_{i,t+\tau} (\sum_{i=1}^{N^*} \alpha_i S_t P^*_{i,t})^{-1} \in \mathscr{G}^f_{t+\tau}$. Define the integration measure, $D(\mathscr{M}, \mathscr{M}^f)$, between the home and the foreign market as the minimal distance between the pricing kernels $m_{t+\tau} \in \mathscr{M}_{t+\tau}$ and $m^f_{t+\tau} \in \mathscr{M}^f_{t+\tau}$, defined on the spaces $\mathscr{G}_{t+\tau}$ and $\mathscr{G}^f_{t+\tau}$, respectively, as:

$$D(\mathscr{M}, \mathscr{M}^f) = \min_{m \in \mathscr{M}, m^f \in \mathscr{M}^f} \| m_{t+\tau} - m^f_{t+\tau} \| \tag{5}$$

where the norm $\|X\| = \sqrt{E[X^2]}$. The measure $D(\mathscr{M}, \mathscr{M}^f)$ captures the degree of mispricing on typical (unit norm) gross returns. This notion of integration is tightly linked with the notion of international risk sharing: when stochastic discount factors between any two countries are equal (not just perfectly correlated), international risk sharing is perfect. Given our set up in terms of gross returns, implying unitary equilibrium price levels, we can interpret this distance $D(\mathscr{M}, \mathscr{M}^f)$ as an upper bound on the percentage pricing differences for portfolios with unit norm gross returns, $ppe(X)$:

$$ppe(X) \leqslant D(\mathscr{M}, \mathscr{M}^f) \quad \text{for } \|X\| = 1 \tag{6}$$

The obtained distance measure can be seen as a measure of the cross-market failure of the law of one price and hence as a measure of imperfect financial integration. The higher the distance measure, the higher will be the percentage pricing difference on portfolios with identical (gross) return characteristics.

DOI: 10.1002/ijfe

Note that the above analysis is done without imposing a positivity constraint on the pricing kernels. We opt for this approach as we are primarily interested in assessing the degree of cross-market integration or segmentation. Irrespective of whether the set of pricing kernels is restricted or not, violations of the law of one price are inconsistent with market integration. In short, for our purpose the weak form integration measure (see CK) is sufficient to measure international cross-market integration.[2]

3. EMPIRICAL RESULTS

In this section we present the results of implementing the integration measures. We find that failures of the law of one price are quite substantial. Subsequently, we analyze possible market characteristics that may explain the existence of this lack of international integration.

3.1. *Data and implementation*

To analyze whether submarkets of the European financial market are integrated or to what extent pricing differentials exist between them, we estimate pair-wise integration measures for a German (DEM) and a US (USD) investor investing in other financial markets being part of the European financial market: Italy (ITL), Spain (ESP), France (FRF), the Netherlands (NLG) and Belgium (BEF). Importantly, these pair-wise integration measures only allow us to draw conclusions on the pair-wise degree of integration. Even if all these markets are pair-wise integrated, the complete European market might still be characterized by segmentation. We also compute an intra-Germany and intra-US integration measure by splitting the German and US markets into two submarkets. These two measures can serve as a benchmark for the other integration measures and allow us to put the cross-market measures more in perspective. In total this gives fourteen pair-wise integration measures.

The data we use are the constituents of Datastream Global Market Indices for the seven markets.[3] We gathered monthly data for the period January 1995 until December 2002. The data are in local currency. End-of-month exchange rates (DEM/euro per foreign currency and USD per foreign currency) are used to convert the payoffs to a common currency. All data are from Datastream.

Each integration measure between any two countries is estimated for 1000 random submarkets of these two countries. Each submarket is constructed by randomly selecting N (N^*) equally weighted portfolios consisting of randomly selected assets from the total number of assets available in each market.[4] As correctly noted by Ayuso and Blanco (2001), the cross-sectional dimension $N + N^*$ should be larger than the time-series dimension T. Otherwise, the system is underidentified and the distance measure always converges to zero. Note also that the integration measure is non-decreasing in the number of base assets. This is quite intuitive, since the more assets are traded, the more demanding it is to maintain pricing consistency among any two markets. Keeping this in mind, each submarket consists of 25 portfolios (of 10 assets).[5] Each portfolio constructed this way is referred to as a base portfolio. Repeating this procedure for 1000 submarkets, yields 1000 integration measures and allows us to draw more robust conclusions on the distance measures. Finally, in computing these integration measures, we account for the impact of the introduction of the euro. This is done by breaking up the time-series in two subsamples (with time-series length $T = 48$). The first sample is the pre-EMU period from January 1995 until December 1998. The second subperiod is the EMU period from January 1999 until December 2002. This gives two symmetrical subsamples of four years around the euro-introduction, and yields insights into the evolution of possible pricing discrepancies.

3.2. *Results*

Table 1 reports the results for the pair-wise integration measures between the sets of pricing kernels for the period 1995–1998. The distance measures are substantial.[6] Comparing to the benchmark distance measures (DEM–DEM or USD–USD), all of the bilateral measures are consistently higher. We test the hypothesis that the population means of the cross-market measures and corresponding benchmark

DOI: 10.1002/ijfe

Table 1. Integration measure (1995–1998)

$D(\mathcal{M}, \mathcal{M}^f)$	Mean	St. Dev.	Min.	Max.
DEM viewpoint				
ITL	0.3376*	0.1748	0.0210	1.1879
ESP	0.3311*	0.1735	0.0177	1.1743
FRF	0.3946*	0.2075	0.0138	1.3432
DEM	0.2468	0.1294	0.0234	0.7731
NLG	0.3666*	0.1989	0.0256	1.1726
BEF	0.5194*	0.2690	0.0238	1.4505
USD	0.3697*	0.1993	0.0333	1.2316
USD viewpoint				
ITL	0.3714*	0.1855	0.0203	1.0328
ESP	0.3588*	0.1900	0.0304	1.1822
FRF	0.3841*	0.2075	0.0331	1.1889
DEM	0.3711*	0.1925	0.0181	1.2108
NLG	0.4118*	0.2158	0.0153	1.3742
BEF	0.4945*	0.2641	0.0448	1.5783
USD	0.3168	0.1746	0.0245	0.9650

The table reports summary statistics of the integration measures (equation (5)) from the viewpoint of a German (DEM) and US (USD) investor. For each market combination, 1000 measures are computed by randomly selecting assets from each market. We report the mean, the standard deviation, the minimum and the maximum of the integration measures. Moreover, we test for equality of group means between the respective cross-market measures and the corresponding benchmark measure (intra-Germany and intra-US). *Indicates that the null hypothesis of equal means is rejected at the 5% significance level ($Z_c = 1.96$). The measures are based on monthly gross returns for the period 1995–1998.

measure, respectively, are equal, and have to reject this hypothesis at the 5% significance level ($Z_c = 1.96$). This indicates larger cross-border price differentials, compared to intra-market price differences. Looking at the DEM viewpoint distance measures, it can be seen that EMU combinations yield, in general, lower distance measures compared to the USD viewpoint distance measures. This indicates that EMU pricing differentials are lower compared to non-EMU pricing differentials for the period 1995–1998. The integration process in EU and the prospect of EMU, then, have led to more similar pricing across countries.

Whether pricing discrepancies reduced with the introduction of the euro can be seen in Table 2. This table reports the results of the integration measures for the period 1999–2002. Looking at the DEM viewpoint we see that monetary unification did not fully eliminate pricing discrepancies. The average values of mispricing in the top panel of Table 2 are still large and significantly different from the benchmark distance measure at the 5% significance level. We do find, however, that pricing differences decreased over time, except for two market combinations, i.e. DEM-ITL and DEM-ESP. For these two market combinations a slight increase in the distance measures can be observed, indicating increased pricing discrepancies. For all of the other markets, there is a decline in the pricing differentials. So, while monetary unification did not fully eliminate the pricing differences, we do observe a decrease in the average mispricing in most markets. Note that the decrease in pricing differences can also be observed for the market combinations from the USD point of view. Moreover, we tend to find a stronger decrease in the mispricing for the USD viewpoint compared to the DEM viewpoint. Apparently, comparing Tables 1 and 2 financial integration is a global trend, as exemplified in the decrease of the average pricing differentials. The hypothesis that the population means of the cross-market measures for the period 1995–1998 and 1999–2002, respectively, are equal, has to be rejected at the 5% significance level for most of the distance measures ($Z_c = 1.96$). This implies that the average decrease in pricing differentials is statistically significant, indicating increased integration. Only for the ITL-DEM and ESP-DEM cases, where we find an increase in the integration measure, the hypothesis of equal means in both time periods cannot be rejected. This means that the increase in average pricing differentials is not statistically significant.

DOI: 10.1002/ijfe

Table 2. Integration measure (1999–2002)

$D(\mathscr{M}, \mathscr{M}^f)$	Mean	St. Dev.	Min.	Max.
DEM viewpoint				
ITL	0.3515*	0.1826	0.0210	1.1862
ESP	0.3459*	0.1826	0.0248	1.0189
FRF	0.2827*,†	0.1474	0.0262	0.8164
DEM	0.2078†	0.1138	0.0173	0.7815
NLG	0.3109*,†	0.1568	0.0169	0.9341
BEF	0.3472*,†	0.1776	0.0119	1.0145
USD	0.2935*,†	0.1515	0.0198	0.9862
USD viewpoint				
ITL	0.3228*,†	0.1725	0.0268	1.1432
ESP	0.3292*,†	0.1794	0.0170	1.1820
FRF	0.2517*,†	0.1279	0.0272	0.7272
DEM	0.3021*,†	0.1585	0.0301	0.9110
NLG	0.2940*,†	0.1529	0.0241	0.9608
BEF	0.3347*,†	0.1751	0.0218	0.9570
USD	0.2295†	0.1214	0.0125	0.7334

The table reports summary statistics of the integration measures (equation (5)) from the viewpoint of a German and US investor. For each market combination, 1000 measures are computed by randomly selecting assets from each market. We report the mean, the standard deviation, the minimum and the maximum of the integration measures. Moreover, we test for equality of group means between the respective cross-market measures and the corresponding benchmark measure (intra-Germany and intra-US). Finally we also test for increased integration over time by testing equality of group means between the respective cross-market measures for 1995–1998 and 1999–2002. The measures are based on monthly gross returns.

*Indicates that the null hypothesis of equal means (between the cross-market measure and benchmark) is rejected at the 5% significance level ($Z_c = 1.96$).

†Indicates that the null hypothesis of equal means (between the 1995–1998 and 1999–2000 measures) is rejected at the 5% significance level ($Z_c = 1.96$).

Whereas integration increased over time, the distance measures indicate that there is still a significant degree of segmentation, implying that large failures of the law of one price are present. This result seems in sharp contrast to the academic belief that financial markets offer no apparent failures of the law of one price (Welch, 2000). In this context, it is important to note that the distance measures reported here constitute an upper bound on pricing differentials. The discount factors are constructed such that they have to price *any* linear combination of the base portfolios. More specifically, the reported distance measures constitute an upper bound to any linear, L^2-integrable, combination of the base portfolios. These linear combinations can, among other things, consist of extreme short selling positions or extreme risk positions in specific base portfolios. As in CK, we computed the portfolio sequences of the pricing difference-maximizing payoffs (CK, equation (31)). We find that 50% of the portfolio weights is negative. In this case, it is not surprising that large mispricings can be detected. The integration measures, thus, crucially depend on the assumption of frictionless markets. In reality, however, several payoff combinations are impossible to construct due to institutional constraints, short selling constraints and transactions costs (bid-ask spreads, margins, commissions). In this sense, mispricing can be due to limits to arbitrage. These limits result from the risky and costly arbitraging process. The degree of mispricing reported here should, therefore, not be confused with textbook arbitrage opportunities. Mispricing equals an arbitrage opportunity only when a sure positive return, with positive probability, can be obtained at *zero cost*. In this respect, the evidence presented here might, indeed, constitute evidence of the existence of important institutional limits to arbitrage. Evidence on institutional limits to arbitrage in closed-end funds, dual listed companies, hedge funds and equity carve-outs is provided by Schleifer and Vishny (1997), Froot and Dabora (1999), de Jong *et al.* (2003) and Lamont and Thaler (2003).

To put these upper bounds on pricing differences somewhat in perspective, we also compute cross-market pricing differences on simple and feasible portfolios. More specifically, we use the unit norm, equally

DOI: 10.1002/ijfe

Table 3. Pricing differentials (1995–1998)

$ppe(X)$	Mean	St. Dev.	Min.	Max.
DEM viewpoint				
ITL	0.0254*	0.0254	1.1E − 06	0.2290
ESP	0.0205*	0.0197	5.4E − 07	0.3131
FRF	0.0189*	0.0186	1.5E − 06	0.1657
DEM	0.0030	0.029	6.2E − 08	0.0276
NLG	0.0134*	0.0151	4.9E − 07	0.1829
BEF	0.0197*	0.0195	4.1E − 07	0.1840
USD	0.0207*	0.0222	1.2E − 06	0.3384
USD viewpoint				
ITL	0.0273*	0.0276	1.9E − 06	0.1999
ESP	0.0259*	0.0236	7.3E − 07	0.1833
FRF	0.0244*	0.0244	1.2E − 06	0.2803
DEM	0.0217*	0.0232	4.8E − 07	0.2477
NLG	0.0215*	0.0245	1.5E − 06	0.3148
BEF	0.0279*	0.0287	2.0E − 06	0.3176
USD	0.0044	0.0042	9.4E − 08	0.0459

The table reports summary statistics of the percentage pricing differences using equation (6) for the viewpoint of a German and US investor. The pricing discrepancies arise when pricing a simple equally weighted foreign portfolio of assets with the home stochastic discount factor for which the minimum distance is reached. We report the mean, the standard deviation, the minimum and the maximum of the percentage pricing difference. Moreover, we test for equality of group means between the respective cross-market measures and the corresponding benchmark measure (intra-Germany and intra-US).
*Indicates that the null hypothesis of equal means is rejected at the 5% significance level ($Z_c = 1.96$).The measures are based on monthly gross returns for the period 1995–1998.

weighted portfolios out of the foreign market as the simple base portfolios. Using equation (6) we calculate the pricing discrepancy that arises if one prices an equivalent portfolio in the home market (with the same stochastic properties in the home market). The computed pricing differences, then, refer to pricing differences of specific and simple equally weighted portfolios (with positive portfolio weights). If extreme positions are to explain the relatively high values of the integration measures, we expect to find relatively small pricing differences for the simple portfolios that we construct. The results of these pricing differences are reported in Tables 3 and 4, for the period 1995–1998 and for the period 1999–2002, respectively. For the period 1995–1998 in Table 3, we see that observed pricing differences are much smaller than suggested by the distance measures. The pricing differentials reported here are small, lower than 3%. Pricing in both markets is very similar, suggesting that, as far as simple portfolios are concerned, there is efficient cross-border pricing. These results are in line with Pagano and Röell (1993) who find that cross-listed companies are equally priced on different stock exchanges. However, we do find that the average percentage pricing difference is higher across markets than within the benchmark market. This means that pricing differences, although small, are still present and create some degree of market segmentation. Table 4 reports the pricing differentials for the EMU-period 1999–2002. We see that pricing differences decreased significantly compared to the pre-EMU period. We have to reject at 5% significance level that the average percentage pricing difference is equal in the 1995–1998 and 1999–2002 period. The only exception is the DEM–DEM benchmark measure. Here, the average is not statistically different in the two time-periods. Moreover, the pricing differences did not only decrease on average, but also the maximum pricing differences decreased substantially. This suggests that markets became more integrated over time. This is certainly true for the DEM viewpoint, suggesting that the introduction of the euro had a positive impact on the elimination of pricing discrepancies. Finally, note that even in the EMU period markets are not yet fully integrated. The cross-border percentage pricing differences are, on average, statistically different from the corresponding benchmark measure, implying still higher pricing differences on an international level compared to the national level.

DOI: 10.1002/ijfe

Table 4. Pricing differentials (1999–2002)

$ppe(X)$	Mean	St. Dev.	Min.	Max.
DEM viewpoint				
ITL	0.0158*,†	0.0166	3.0E − 07	0.2142
ESP	0.0144*,†	0.0133	3.1E − 06	0.0883
FRF	0.0121*,†	0.0121	5.4E − 07	0.1316
DEM	0.0029	0.0030	5.6E − 08	0.0406
NLG	0.0112*,†	0.0105	3.9E − 08	0.0966
BEF	0.0128*,†	0.0144	1.4E − 07	0.1737
USD	0.0174*,†	0.0156	5.3E − 07	0.1123
USD viewpoint				
ITL	0.0244*,†	0.0233	9.1E − 07	0.2453
ESP	0.0237*,†	0.0246	1.9E − 07	0.2158
FRF	0.0174*,†	0.0169	1.5E − 07	0.2066
DEM	0.0177*,†	0.0171	1.7E − 06	0.1672
NLG	0.0188*,†	0.0183	2.0E − 06	0.2049
BEF	0.0226*,†	0.0214	2.7 − 07	0.2389
USD	0.0040†	0.0039	2.4E − 07	0.0455

The table reports summary statistics of the percentage pricing differences using equation (6) for the viewpoint of a German and US investor. The pricing discrepancies arise when pricing a simple equally weighted foreign portfolio of assets with the home stochastic discount factor for which the minimum distance is reached. We report the mean, the standard deviation, the minimum and the maximum of the percentage pricing differences. Moreover, we test for equality of group means between the respective cross-market measures and the corresponding benchmark measure (intra-Germany and intra-US). Finally, we also test for increased integration over time by testing equality of group means between the respective pricing discrepancies for 1995–1998 and 1999–2002. The measures are based on monthly gross returns.

*Indicates that the null hypothesis of equal means (between the cross-market and benchmark pricing differential) is rejected at the 5% significance level ($Z_c = 1.96$).

†Indicates that the null hypothesis of equal means (between the 1995–1998 and 1999–2000 pricing differentials) is rejected at the 5% significance level ($Z_c = 1.96$).

3.3. Market structure and arbitrage

The integration analysis indicates that the degree of financial integration, as measured by the upper bound on pricing discrepancies varies both across the market combination analyzed, as well as across time. In general, we find a decreasing trend in pricing differentials, indicating increased financial integration. Yet, quite some cross-sectional variation across market combinations in the pricing discrepancies persists. In this section, we study factors that may explain both the time variation as well as the cross-sectional variation in financial integration. Motivated by the micro-finance literature, we focus on market characteristics as a possible explanation for the observed limits to arbitrage. The theory of market microstructure is built on the idea that asset prices might depart from its rational expectations value due to a variety of frictions (see Madhavan, 2000). These frictions can be real or informational (Stoll, 2000). The former are due to processing costs, inventory risk or monopoly power, while the latter can be caused by the free trading option which is offered by quotes as well as by the presence of asymmetric information. These trading frictions make it more difficult to trade assets, and hence can be expected to create a degree of mispricing. In the market microstructure literature the bid-ask spread is a commonly used measure of friction. A wide bid-ask spread causes the prices at which assets are traded to deviate from their rational expectations value. This idea is elaborated in this section.

The methodology used here is in the same spirit as the Demsetz (1968) model who studies the determinants of the spread between the bid and the ask price. In line with Demsetz (1968), we model the spread between two stochastic discount factors (for which the minimal distance is reached) as a function of market value, volatility, market activity and an integration trend. The rationale for these variables is primarily based on processing costs and inventory risk. Market value should yield a negative coefficient. The price impact of block trades has shown to be large in small market capitalizations (see Keim and Madhavan, 1996). When market value is considerable, large trades have no significant impact on the

DOI: 10.1002/ijfe

market. When market value is small, large trades might have an impact on the market, creating distortions in the pricing process. Also, pricing of assets is more efficient and less costly in a large market compared to a small market, since the probability of finding a counterparty is higher in the former. The coefficient of market volatility should be positive. Volatility, typically created by noise traders, induces departures from the fundamental values (see De Long *et al.*, 1990). A higher volatility is accompanied by larger pricing departures, and thus, larger pricing differences. Market activity, defined as the market turnover divided by the market value, should have an inverse relation with pricing differences. This factor accounts for a liquidity premium (see Brennan *et al.*, 1998). The higher the market activity, the higher the liquidity, and thus the lower the liquidity premium and pricing differences. Finally, financial markets are becoming more and more global and transparent, lowering cross-border pricing differentials. This process is captured by the inclusion of a time-trend representing increased integration over time.

We estimate the following panel:

$$|m_t - m_{t,i}| = \alpha_0 + \alpha_1 dMV_{t-1,i} + \alpha_2 \sigma_{t-1,i} + \alpha_3 MA_{t-1,i} + \alpha_5 TREND_t + \varepsilon_{t,i} \tag{7}$$

with m_t the stochastic discount factor of the German market and $m_{t,i}$ the stochastic discount factor of the country i, for $i = ITL$, ESP, FRF, DEM, NLG, BEF and USD and $t = 1995.1, \ldots, 2002.12$. The variables $dMV_{t,i}$, $\sigma_{t,i}$, $MA_{t,i}$ and $TREND_t$ represent the change in the market value (relative to its long term trend), the market volatility, the market activity and the time-trend, respectively.[7] We estimate a panel such that the above model is estimated jointly for all countries i. A similar panel is estimated for the USD viewpoint, with $i = ITL$, ESP, FRF, DEM, NLG, BEF and USD.

The difference between the stochastic discount factors used as the dependent variable results from the integration measures. For each of the 1000 pair-wise measures we compute the absolute difference between the stochastic discount factors for which a minimum is reached. Subsequently, we average over these 1000 absolute differences to yield a single time-series of absolute differences between the stochastic discount factors for each market combination. The data of the market characteristics are monthly data from the Datastream Global Market Indices for the seven countries. Market value is the share price multiplied by the number of shares in issue. However, since market value contains a unit root, we look at the percentage change in market value relative to its long-term trend, making use of the Hodrick–Prescott filter. Monthly volatilities are computed as the standard deviations based on daily price data. Finally, market activity is computed as the market turnover, divided by the market value. The data of these factors are from Datastream. The summary statistics of the market properties are reported in Table A1 in Appendix A. We see that there is substantial cross-variation in the factors. The largest absolute cross-variation can be found for the market activity indicator. Market activity data show that Belgium is the least active market, while Germany is the most active market. A similar difference can be observed for the change in market value, with the smallest change in market capitalization for Belgium and the largest change for the US. Finally, Belgium is also the least volatile market and Italy the most volatile market.

The results of the two panel estimations (7) are reported in Table 5. All of the factors are estimated significantly and the signs of the estimated coefficients are as expected. The effect of market capitalization (dMV) is negative. A positive market growth corresponds to lower pricing differences. Also, volatility (σ) has a positive estimate. A more volatile market is characterized by more pricing discrepancies. Furthermore, an increase in market trading activity (MA) induces lower pricing differences. Finally, the integration dummy has a negative estimate on the pricing differentials. Pricing differences declined over time. This represents increased global financial integration. The R^2 of the USD panel is rather large. The market factors explain almost 25% of the variance in the pricing differences. The explanatory power of the market factors in the DEM panel is much smaller: the market factors only explain 16% of the variance in the pricing differentials. Finally, the F-statistic tests the hypothesis that all estimates are jointly zero. This test is strongly rejected for both panels (5% critical value of the $F_{4,654} = 2.60$). Market properties are, thus, able to explain a considerable part of the cross-border pricing differences. This holds especially for the USD panel. A detailed analysis of the regression results and the fits do show that the factors in the regression framework (more in particular all independent variables but the time trend) do fit quite well the cross-sectional variation in the pricing discrepancies. The time series properties of the pricing discrepancies are only fit by the time trend.

DOI: 10.1002/ijfe

Table 5. Pricing differentials explained

Parameter	Estimate	St. error	p-value
DEM viewpoint			
α_0	0.3043	0.0073	0.000
$dMV_{t-1,i}$	−0.1173	0.0232	0.000
$\sigma_{t-1,i}$	1.6703	0.5265	0.002
$MA_{t-1,i}$	−0.4246	0.0516	0.000
$TREND_t$	−0.0006	0.0001	0.000
$R^2 = 16.1\%$ $F = 30.946$			
USD viewpoint			
α_0	0.3129	0.0064	0.000
$dMV_{t-1,i}$	−0.0651	0.0255	0.011
$\sigma_{t-1,i}$	2.9435	0.5182	0.000
$MA_{t-1,i}$	−0.2199	0.0448	0.000
$TREND_t$	−0.0013	0.0001	0.000
$R^2 = 24.9\%$ $F = 53.60$			

Estimate gives the estimated coefficients of the panel given in (7), St. error and p-value are the corresponding standard error and p-value of the estimate, respectively. We also report the model fit (R^2) and the F-statistic testing the hypothesis that all coefficients are jointly zero. The 5% critical value for the $F_{4,654} = 2.60$. The model is estimated first from the viewpoint of a DEM investor, next from the viewpoint of a USD investor.

Table 6. Contribution of the micro-factors in the fit of the model

	Model fit	α_0	$\alpha_1 dMV$	$\alpha_2\sigma$	$\alpha_3 MA$	$TREND$
DEM viewpoint						
ITL	0.2698	0.3043	−0.0001	0.0229	−0.0291	−0.0282
ESP	0.2681	0.3043	−0.0002	0.0200	−0.0278	−0.0282
FRF	0.2710	0.3043	−0.0002	0.0202	−0.0251	−0.0282
DEM	0.2445	0.3043	0.0002	0.0192	−0.0510	−0.0282
NLG	0.2519	0.3043	0.0001	0.0185	−0.0428	−0.0282
BEF	0.2823	0.3043	0.0001	0.0144	−0.0083	−0.0282
USD	0.2532	0.3043	−0.0001	0.0224	−0.0452	−0.0282
USD viewpoint						
ITL	0.2748	0.3129	3.9E − 05	0.0380	−0.0150	−0.0611
ESP	0.2723	0.3129	9.8E − 05	0.0347	−0.0143	−0.0611
FRF	0.2728	0.3129	0.0001	0.0339	−0.0130	−0.0611
DEM	0.2597	0.3129	0.0003	0.0339	−0.0263	−0.0611
NLG	0.2620	0.3129	0.0002	0.0321	−0.0221	−0.0611
BEF	0.2760	0.3129	0.0002	0.0283	−0.0043	−0.0611
USD	0.2608	0.3129	0.0002	0.0321	−0.0233	−0.0611

The table reports the contribution to the models' fit by the different micro-factors. This contribution of the different micro-factors is computed for their average values reported in Table A1 in Appendix A.

In order to put the different effects more in perspective, we compute the average contribution of the different market factors to the fit of the stochastic discount factor differences. That is, we compute the contribution of the different factors for the average values of the independent variables reported in Table A1 (Appendix A). Table 6 reports these results. From the table it is clear that market activity is the most important factor in the determination of the pricing differences. The higher the market activity, the smaller the pricing differentials. This effect is most clear for the DEM viewpoint, where we see a larger cross-sectional variation in the effect of market activity. Also market volatility is an important factor: the larger the volatility on a certain market, the higher the pricing differences will be. Finally, the time trend is an important factor, whereas the change in market capitalization is the least important factor.

DOI: 10.1002/ijfe

Next to decomposing the average absolute pricing errors as done in Table 6, one can also gauge the importance of each of the independent variables from a time-series perspective. Multiplying the estimated (semi-) elasticities by the average size of the shocks to the independent variables (i.e. the standard deviations reported in Table A1) yields the average time variation in the absolute pricing differentials. Typically, we find in this dimension that market value dominates. For instance, computing absolute changes in the pricing differentials (DEM viewpoint) with respect to typical shocks yields 0.0117, 0.0083 and 0.0085 for market capitalization, volatility and market activity, respectively.[8] So while the cross-section is dominated by differentials in market activity, time variability of the bound is more responsive to a change in market capitalization.

4. CONCLUSION

In this paper, we analyzed the degree of international financial integration across some of the major financial markets in the world. We used the methodology of Chen and Knez (1995) to measure financial integration by the law of one price. Subsequently, we related the recovered arbitrage pricing differentials to market characteristics. From a methodological point of view this study adds to the literature by extending the set of assets on which relative pricing relations can be tested. In principle we can extend the set of assets to all traded assets. This extension of the cross-sectional sample size is important as it allows us to apply standard econometric tools to study the interaction between the market characteristics of financial markets and the degree of mispricing.

The main conclusions to draw from this analysis are twofold. First, unlike the literature on cross-listings, we find relative pricing differentials, i.e. failures of the law of one price, to be quite substantial, even for the most developed financial markets. The main reason for this finding is that, using the CK metric, one assesses all types of possible failures of the law of one price and not only the most apparent ones, such as on the cross-listed assets. Even though there is convincing evidence that cross-listed assets are priced according to the law of one price, our results suggest that there exist (equivalent) portfolios across markets for which substantial mispricing is observed. These portfolios may, however, be very complex, imply extreme short positions and be, in practical terms, infeasible. The second finding of this study is that the cross-market mispricings are to some extent explained by financial market characteristics. We find that standard characteristics of financial markets do explain a substantial part of the observed pricing discrepancies. More specifically, the change in market capitalization, market volatility and market activity are important determinants of the observed mispricings. These factors appear both statistically and economically important and square well with the explanation set out in the micro-finance literature (see Demsetz, 1968).

APPENDIX A: DATA DESCRIPTION

The summary statistics of the market properties are reported in Table A1.

Table A1. Summary statistics market characteristics

		Mean	St. Dev.	Min.	Max.
DEM viewpoint					
Change in market capitalization (%)	ITL	0.0082	0.1631	−0.2277	0.7183
	ESP	0.0016	0.1100	−0.2270	0.3371
	FRF	0.0013	0.1190	−0.2691	0.2864
	DEM	−0.0019	0.1110	−0.2770	−0.3487
	NLG	−0.0011	0.0926	−0.2543	0.2432

DOI: 10.1002/ijfe

Table A1. (*continued*)

		Mean	St. Dev.	Min.	Max.
	BEF	−0.0010	0.0996	−0.1901	0.2933
	USD	0.0080	0.1079	−0.2263	0.3390
Market volatility	ITL	0.0137	0.0056	0.0067	0.0356
	ESP	0.0120	0.0049	0.0051	0.0300
	FRF	0.0121	0.0052	0.0049	0.0298
	DEM	0.0115	0.0055	0.0034	0.0297
	NLG	0.0111	0.0064	0.0038	0.0330
	BEF	0.0086	0.0049	0.0028	0.0281
	USD	0.0134	0.0050	0.0061	0.0284
Market activity	ITL	0.0685	0.0254	0.0235	0.1268
	ESP	0.0654	0.0188	0.0274	0.1076
	FRF	0.0591	0.0148	0.0270	0.1085
	DEM	0.1200	0.0906	0.0025	0.3575
	NLG	0.1009	0.0337	0.0506	0.2048
	BEF	0.0195	0.0042	0.0114	0.0319
	USD	0.1064	0.0294	0.0600	0.1815
USD viewpoint					
Change in market capitalization (%)	ITL	−0.0006	0.1321	−0.2374	0.3365
	ESP	−0.0015	0.0853	−0.1764	0.2069
	FRF	−0.0020	0.0911	−0.2199	0.1924
	DEM	−0.0041	0.0863	−0.2309	0.1808
	NLG	−0.0024	0.0638	−0.1991	0.1234
	BEF	−0.0037	0.0940	−0.1435	0.3092
	USD	−0.0025	0.0710	−0.1807	0.1794
Market volatility	ITL	0.0129	0.0051	0.0057	0.0345
	ESP	0.0118	0.0047	0.0044	0.0264
	FRF	0.0115	0.0049	0.0044	0.0291
	DEM	0.0115	0.0050	0.0040	0.0268
	NLG	0.0109	0.0055	0.0039	0.0324
	BEF	0.0096	0.0044	0.0036	0.0286
	USD	0.0109	0.0049	0.0032	0.0256

The table shows the summary statistics of the % change in market capitalization (dMV), of market volatility (σ) and of market activity (MA). Since market capitalization and market volatility are viewpoint dependent, they are given both in DEM and in USD currency. Market activity is viewpoint independent. Data are on a monthly basis for the period 1995–2002.

NOTES

1. More specifically, the only assumption to be made is the assumption of mean square integrable pricing processes.
2. This implies that integration is studied by the law of one price, and not by the absence of arbitrage opportunities. In general, the inclusion of absence of arbitrage opportunities, i.e. imposing a positivity constraint on the stochastic discount factors, makes the bound on stochastic discount factors tighter, and thus the integration metric higher.
3. The Datastream Global Market Indices contain the most representative stocks in a country. From the total sample, we selected the stocks that are included in the index for the complete time-series period of January 1995 until December 2002. For Germany this gives 143 assets, for the US 765 assets, for Italy 88 assets, for Spain 75 assets, for the Netherlands 92 assets, for France 148 assets and for Belgium 53 assets.
4. CK note that using portfolios instead of individual assets reduces measurement errors.
5. Since this procedure is repeated 1000 times, we checked whether it did not cause an important overlap in the sample studied. We found that the probability of comparing two equal submarkets is extremely small (10 000 random draws did not yield a single same market combination). Finally, note also that the 25 base portfolios forming a submarket are always different.
6. Recall, however, that equal pricing kernels as defined in the integration measure, crucially depend on the assumption of market completeness. Therefore, the computed degree of segmentation might also be attributed (partly) to market incompleteness.
7. To avoid the problem of endogeneity, we consider one-period lagged values of the explanatory variables.
8. Typical shocks are chosen at 0.1, 0.005 and 0.02 for market capitalization, volatility and market activity, respectively.

DOI: 10.1002/ijfe

REFERENCES

Ayuso J, Blanco R. 2001. Has financial integration increased during the 1990s? *Journal of International Financial Markets*, *Institutions and Money* **11**: 265–287.

Backus DK, Foresi S, Telmer CI. 2001. Affine term structure models and the forward premium anomaly. *Journal of Finance* **56**: 279–304.

Bodurtha J, Kim DS, Lee C. 1995. Closed-end country funds and the US market sentiment. *Review of Financial Studies* **8**: 879–918.

Brandt MW, Cochrane JH, Santa Clara P. 2004. International risk sharing is better than you think (or exchange rates are much too smooth). Working Paper, UCLA.

Brandt MW, Santa-Clara P. 2002. Simulated likelihood estimation of diffusions with an application to exchange rate dynamics in incomplete markets. *Journal of Financial Economics* **63**: 161–210.

Brennan MJ, Chordia T, Subrahmanyam A. 1998. Alternative factor specifications, security characteristics, and the cross-section of expected stock returns. *Journal of Financial Economics* **49**: 345–373.

Chen Z, Knez PJ. 1995. Measurement of market integration and arbitrage. *Review of Financial Studies* **8**: 287–325.

de Jong A, Rosenthal L, van Dijk MA. 2003. The limits of arbitrage: evidence from dual-listed companies. Mimeo.

de Long JB, Schleifer A, Summers LH, Waldmann RJ. 1990. Noise trader risk in financial markets. *Journal of Political Economy* **98**: 703–738.

Demsetz H. 1968. The cost of transacting. *Quarterly Journal of Economics* **82**: 33–53.

Froot KA, Dabora E. 1999. How are stock prices affected by the location of trade? *Journal of Financial Economics* **53**: 189–216.

Hansen LP, Jagannathan R. 1997. Assessing specification errors in stochastic discount factor models. *Journal of Finance* **52**: 557–590.

Hardouvelis G, La Porta R, Wizman T. 1994. What moves the discount on country equity funds? In *The Internationalization of Equity Markets*, Frankel JA (ed.). University of Chicago Press: Chicago.

Keim DB, Madhavan A. 1996. The upstairs market for large-block transactions: analysis and measurement of price effects. *Review of Financial Studies* **9**: 1–36.

Lamont OA, Thaler RH. 2003. Can the market add and subtract? Mispricing in tech stock carve-outs. *Journal of Political Economy* **111**: 227–269.

Madhavan A. 2000. Market microstructure: a survey. *Journal of Financial Markets* **3**: 205–258.

Pagano M, Röell A. 1993. Shifting gears: an economic evaluation of the reform of the Paris Bourse. In *Financial Markets Liberalization and the Role of the Banks*, Conti V, Hamamiu R (eds). Cambridge University Press: Cambridge.

Schleifer A, Vishny RW. 1997. The limits of arbitrage. *Journal of Finance* **52**: 35–55.

Stoll HR. 2000. Friction. *Journal of Finance* **55**: 1479–1514.

Welch I. 2000. Views of the financial economists on the equity premium and on professional controversies. *Journal of Business* **73**: 501–537.

DOI: 10.1002/ijfe