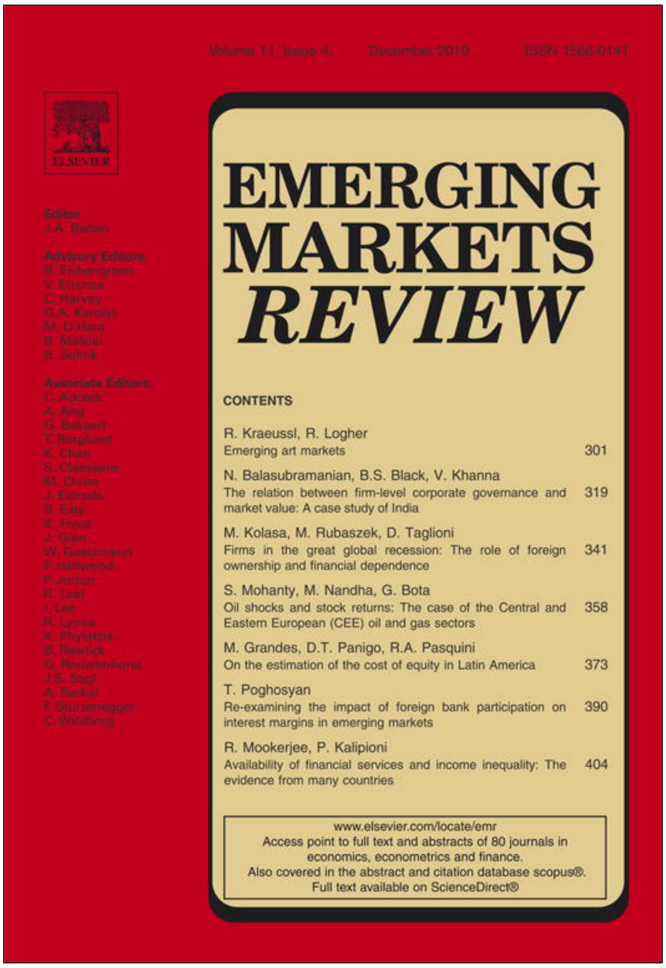
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On the estimation of the cost of equity in Latin America

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This paper researches the sources of stock market risk influencing the pricing of 921 Latin American stocks and computes their corresponding opportunity cost (COE) over the period 1997–2004 by firm and sector. Running an adjusted version of the Capital Asset Pricing Model (CAPM) it finds that systematic risk accounts on average for more than 32% of COE total variance. This implies that potential CAPM mispricing related to undiversified idiosyncratic risk in Latin America has been relatively lower (but absolutely higher) than in United States and other European and Asian stock markets (such as the United Kingdom, Canada or Japan). A first robustness test for the omission of international sources of un-diversifiable risk suggests that both global market and real currencies portfolios do not add significant information to domestic market portfolios. Moreover, a second robustness check offers further evidence that well-diversified portfolios constructed by sorting stocks according to their size and book-to-market ratios *a la* Fama and French do not improve the goodness of fit in the regressions based on the adjusted version of CAPM.

1. Introduction

The weighted average cost of capital (WACC) – a combination of equity and debt costs paid by either public or private entities – is an important determinant of corporate investment decisions and economic growth. In

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Latin America, WACC, and in particular the opportunity cost of equity capital (henceforth COE), has been generally relatively higher and volatile compared to OECD countries excluding Mexico (see Fama and French, 1998; Hail and Leuz, 2004 or Mishra and O'Brien, 2005). This may be attributable to a number of reasons beyond – though not unrelated to – historically weak macroeconomic policy frameworks, macroeconomic volatility, external vulnerability, low national savings and investment rates and as a consequence a remarkable dependence on external financing.

In spite of the reforms and capital market liberalizations of the late 80s and early 90 s, Latin American stock markets are still relatively underdeveloped compared to its East Asian peers (see de la Torre et al., 2008). Indeed, they continue to be characterized by: a) low market capitalization to GDP ratios, low turnover, limited number of publicly traded firms and scant free floats, b) illiquidity (low traded volumes and significant number of days without any transaction) with the exception of Chile, c) high average stock returns, d) a growing number of company stock delistings or cross-listings to achieve lower capital cost and d) stock return volatility in an environment where local interest rates are volatile by international standards and bear positive sovereign default risk with varying levels across countries.

Against this backdrop raises the question of which factors drive COE in Latin America and to what extent the stock excess return and its associated volatility can be accounted for by domestic, global or other sources of risk once we control for illiquidity, macroeconomic instability or the presence of sovereign default risk in local interest rates; or whether idiosyncratic risk matters at all and should be priced in the underlying model. This paper offers new empirical evidence on the sources of risk that determine the opportunity cost of equity capital using seven Latin American stock markets as a case study. We pick the seven largest markets by capitalization, namely Argentina, Brazil, Chile, Colombia, Mexico, Peru and Venezuela spanning monthly observations for the period 1997–2004 but using information since 1993. Building on stock price data from 921 publicly listed firms we are able to design and test a robust time series econometric methodology to estimate COE across countries, sectors and firms.

The study makes a contribution to the empirical literature on stock pricing and corporate valuation in emerging markets (see Godfrey and Espinosa, 1996; Rouwenhorst, 1999; Wong, 2000; Fama and French, 1998, 2004; Bekaert and Harvey, 2000; Bonomo and Garcia, 2001; Bruner et al., 2002; Mishra and O'Brien, 2005; Pereiro, 2006; Iqbal et al., 2010; or Barclay et al., 2010 among others) in at least 3 ways:

1) it sets out and econometrically runs a modified version of CAPM to allow for the instability in the estimation of betas — fact not specific to Latin America as evidenced in Brealey et al. (2009), the presence of sovereign risk as government risk-free rates in most of these countries are not default risk-free, something not taken into account in previous literature as the term structure of risk-free rates is assumed not to be affected by sovereign risk or to remain flat over time or a political risk factor is included separately as in Mishra and O'Brien (2005), illiquidity[3] and the stochastic nature of risk-free interest rates,
2) it provides two robustness tests to check whether well-diversified global and multilateral currency portfolios on the one hand, and portfolios constructed according to the stock size and book-to-market ratios, add information to the adjusted version of CAPM on the other, and
3) it informs business and financial managers on how (or how not) to value stocks in a large and representative sample of emerging Latin American markets employing that modified version of CAPM (see for example, Bruner et al., 2002; or Pereiro, 2006, for the case of Argentina).

Taking as starting point an adjusted version of the Capital Asset Pricing Model (CAPM, Sharpe, 1964; Lintner, 1965), we first investigate how much of COE variability can be attributed to systematic (i.e., undiversfiable) risk. We find that for Latin America as a whole, nearly a third of COE variability on average can be attributed to systematic sources of risk. In a second step, we check the robustness of these results to the inclusion of global stock and real currency portfolios (global factors in what follows). In other words, should internationally well-diversified portfolios be factored in the pricing of Latin American stocks? In fact, as Latin American countries opened up its capital markets to foreign investors and let its residents invest abroad late in the 80s or early in the 90s (see Bekaert and Harvey, 2000, or Henry, 2000), the question of how much of the stock risk that Latin American firms bear is correlated with global risk factors has increasingly gained ground in

[3] By weighing each stock return by the number of traded shares at each day with observations. This contrasts with other studies where liquidity is a conditioning factor in the stock pricing equation (see for instance, Liu, 2006; Barclay et al., 2010; or Hearn and Piesse, 2009).

the research agenda. Our results, drew from data corresponding to the post-equity market liberalization period in Latin America suggest that global factors do not add explanatory power to domestic portfolios in the adjusted CAPM regressions, which may signal that a significant fraction of regional/country-specific stock market risk is indeed correlated with global risk and hence that local market portfolios are already capturing the relevant sources of global risk. They also indicate that increasing but still modest correlations between developed country and Latin American stock excess returns (except for Mexico, whose correlation with US stocks is somewhat higher) may reduce the informational power of global portfolios and the diversification benefits of the latter from the domestic investor perspective (Bekaert and Harvey, 2003).

In a third and last step, we extend the econometric analysis over the same sample of stocks to test for the presence of other sources of common risk originated in the size – small firms are required a higher return than larger firms – and book-to market value of stocks that pervade in stock returns and that domestic market portfolios fail to price in according to the Fama–French Three Factor Model (FF3FM, Fama and French, 1992, 1993, 1996, 1998). We conclude that: 1) both size and value premia are not generally statistically significant risk factors, and 2) they do not add informational power to the domestic market portfolio in the explanation of stock (excess) returns.

The remainder of the paper is organized as follows. Section 2 presents the baseline theoretical framework, which is an adjusted version of CAPM, the data, and the econometric estimates of COE across markets, sectors and firms. In the same section we calculate and compare two measures of the share of systematic risk in COE. Section 3 provides 2 robustness tests for the inclusion of additional sources of global factors and common risk not allowed for in CAPM. Section 4 summarizes the results and concludes.

2. COE estimation and the role of systematic risk

2.1. Starting point: Adjusting the standard CAPM

In this section we present an adjusted version of the Capital Asset Pricing Model (CAPM) that takes into account the likely instability of the estimated beta parameters (or systematic risk factor loading), the choice of the risk-free rate under sovereign default risk, the stochastic nature of interest rates and the low frequency trading in emerging stock markets, i.e. illiquidity.

We start from the CAPM version of Sharpe (1964) Lintner (1965). The pricing relation implied by this model yields the opportunity COE, defined as the expected return of an asset (stock) *i*. Eq. (1) illustrates the CAPM pricing relationship:

$$E(R_i) - R_f = \beta_{iM}\left[E(R_M) - R_f\right] \ \forall i = 1, \ldots, N \tag{1}$$

where $E(R_i)$ is the expected return of asset *i*, R_f is the risk-free interest rate, $(E(R_M) - R_f)$ is the domestic market risk premium (or domestic market excess return) and β_{iM} is the individual excess return sensitivity to domestic market risk or the systematic risk factor loading.

Since it is assumed that investors are able to fully diversify their portfolios, all individual risk is assumed to be diversified away so the expected return of an asset i is only determined by the sensitivity to systematic risk, captured by the multiple of beta and the excess return on the market portfolio.

Before presenting our data and estimating Eq. (1) several issues need to be addressed. A first issue concerns the rate which should be used as a risk-free interest rate. While in developed markets the yield on a government bond may be a good proxy for a risk-free rate, in most Latin American markets government/sovereign bond yields do not come without default risk, therefore a spread should be added to the pure risk-free rate. Moreover, sovereign default risk is typically a significant determinant of corporate default risk,[4] and hence accounts for part of the non-diversifiable risk prevailing in those markets. As a result, we proceed to incorporate sovereign default risk into the COE pricing equation, assuming this risk is actually priced in the assets composing the local market portfolio, as follows:

$$COE_{i,t} = R_{f,t} + \beta_i^b\left[R_{M,t} - R_{f,t} - SS_{j,t}\right] + SS_{j,t} \tag{2}$$

[4] See for instance Durbin and Ng (2005). The reader should note that as of the time of writing this article only Chile and Mexico´s government bonds were rated in the investment grade category.

where $R_{f,t}$ is the US Treasury Bond yield, $SS_{j,t}$ is the sovereign spread of country j at time t, $(R_{M,t} - R_{f,t} - SS_{j,t})$ is the sovereign spread-adjusted market excess return, and β_i^b is the asset i excess return sensitivity to the sovereign spread-adjusted market excess return.

A second issue pertains to the choice of the proxy for the theoretical local market portfolio and whether global factors should be incorporated into it. Theoretically, the chosen proxy should completely capture the un-diversifiable risk prevailing in the market. A reasonable choice might be to use a local market index tracking the most representative stocks in terms of trading and capitalization.

A third issue is the observed volatility in interest rates $R_{f,t}$ and the potential instability of the beta estimates. In this regard, our econometric strategy is to estimate an adjusted version of Eq. (2) following Black's (1972) approach to interest rate volatility[5] using monthly stock return observations, a four-year rolling-window[6] sample to control for beta instability (i.e. we estimate moving betas) and the unconditional Generalized Method of Moments (GMM) estimator. We favour GMM against OLS estimators because as it has been documented in the literature which studies the behavior of financial assets returns (see for instance Bonomo and Garcia, 2001, who also use GMM) the OLS estimator suffers from several shortcomings: it has a mean reversion tendency, it is inefficient when returns (or excess returns) distributions are abnormal, and also introduces significant biases when stocks are poorly and thinly traded (illiquidity). On the contrary, the GMM estimator does not rely on the normality, homoskedasticity, and the lack of serial correlation assumptions under OLS. Indeed, GMM estimators can afford distribution of returns which are heteroskedastic, serially dependent and non-Gaussian. The only required assumption is that returns are stationary with finite fourth moments (see Ferson and Jagannathan, 1996; Fernandes, 2004).[7]

A fourth issue has to do with the presence of infrequent and low volume trading in the sample that may bias our estimations. To avoid the loss of observations that would result if we reduced the sample frequency, we use a weighting matrix in our regressions to correct for low-trading observations.[8] The matrix weights are the number of monthly traded shares. In this way, high traded volume drive beta estimates. Finally, to address the potentially high variability of beta estimates we adopt the adjustment procedure put forward by Vasicek (1973) and apply it to our 4-year estimated rolling betas. Individual stock illiquidity, changes in the market indexes (owing to delisting, mergers and acquisitions, etc.) and macroeconomic uncertainty are possible factors sparking highly volatile CAPM beta estimates. Basically, the Vasicek's estimator preserves the mean reversion property while taking into account the dispersion in rolling individual betas and across sectors.[9]

[5] We use Black's (1972) estimation procedure which allows for interest rate variability within the sample period considered, a useful feature in markets that display significant interest rate volatility This is: $Ex_{i,t} = \beta_i^b [R_{M,t} - R_{f,t} - SS_{j,t}] + v_{i,t}$; where $Ex_{i,t}$ stands for "asset excess return", namely: $Ex_{i,t} = R_{i,t} - R_{f,t} + SS_{j,t}$.

[6] Taking into account a potential trade-off between efficiency and the likelihood of structural breaks (i.e. parameter instability), Altman et al. (1974), Roenfeldt et al. (1978), Smith (1980), and Daves et al. (2000) conclude that the optimal estimation period ranges from four to nine years. A four-year sample is consistent with the average length of the business cycle in the largest Latin American countries (see Carrera et al., 1998).

[7] Latin American stock market returns have very small first order autocorrelation coefficients and none of them appear to be close to the unit root (for lack of space we do not report these results here). The results of the Augmented Dickey–Fuller (ADF) tests (without structural breaks in order to allow for a higher probability of non-rejection of the unit root hypothesis) do not support the null hypothesis of the existence of a unit root. Indeed, 90 to 95% of these results (depending on whether we use 12 or 4 lags in the ADF test) reject the null hypothesis as GMM assumptions prescribe. Lastly, given the lack of a widely-accepted underlying theory supporting a conditional GMM equation (a GMM equation with instrumental variables for local market excess returns), we use the unconditional equation yielding GMM estimators that are the same as the OLS estimators. Nevertheless, GMM regressions yield heteroskedasticity and serial correlation consistent standard deviations, which is particularly important for specification purposes.

[8] This is equivalent to assume that investors form expectations using a simple and single rule: factor loadings on systematic risk (beta parameters) are better represented by high-trading observations. If we equally weighed high and low-trading (including non-trading) observations we would introduce noise in the beta expectation formation because we would have a higher probability of hazardous returns in low-trading days.

[9] When the standard deviation of the estimate of the individual beta is high compared to the standard deviation of the average beta estimate, the beta forecast tends to converge to the average beta (e.g. average beta by sector). Conversely, when the standard deviation of the estimate of the individual beta is small compared to the standard deviation in the estimate of the average beta, then the beta forecast tends to converge towards the individual beta. Formally, Vasicek's (1973) adjusted beta is a weighted average between the individual and sector betas where beta standard deviations (the measure of the uncertainty about betas) are the weights:

$$\beta_i^{VA} = \beta_i \left(\frac{\sigma_{\beta_s}}{\sigma_{\beta_s} + \sigma_{\beta_i}} \right) + \beta_s \left(\frac{\sigma_{\beta_i}}{\sigma_{\beta_s} + \sigma_{\beta_i}} \right);$$

where β_i^{VA} is the firm i's Vasicek's adjusted beta, β_i is the firm i's CAPM standard beta, β_s is the sector s's across-firm average beta (with firm i belonging to sector s), $\sigma_{\beta i}$ is the firm i's standard deviation of beta estimate and σ_{β_s} is sector s's across-firm standard deviation of betas.

Summing up, we modify the standard CAPM in order to make it suitable for emerging stock market data by dealing altogether with multiple biases caused by illiquidity, potential instability in the beta parameter estimates and highly volatile interest rates in economic environments where risk-free rates are not risk-free. Assuming that the individual premium on systematic risk already prices in the extra premium due to sovereign risk (i.e., in the case that the sovereign premium is incorporated as part of the excess market portfolio return), the unbiased COE estimation can be performed through the following equation:

$$COE_{i,t} = R_{f,t} + \beta_i^{va}\left[R_{M,t} - R_{f,t} - SS_{j,t}\right] + SS_{j,t} \tag{3}$$

where $SS_{j,t}$ is the local government yield in excess of the corresponding US Treasury bond rate, or sovereign spread.[10] Eq. (3) is the adjusted version of the CAPM equation we adopt to obtain the empirical results reported hereinafter.

2.2. Data

Our dataset covers 921 publicly traded firms from the 7 largest Latin American stock markets: Argentina, Brazil, Chile, Colombia, Mexico, Peru and Venezuela. The main source is *Economatica*. Although the full, raw sample spans the period 1986–2004 we limit the analysis to a restricted sample from 1993 to 2004 in order to get comparable figures. However, as the econometric models are run on the basis of 48-month moving windows our COE estimates start in 1997 using information since 1993. All data are expressed in current US dollars.

Table 1 presents summary statistics about the seven stock markets. Column I displays the absolute number of firms (stocks) per country reported by *Economatica* and Column II indicates the percentage share of these firms in the total number of companies listed in each stock market at the end of 2004 according to the World Federation of Stock Exchanges. Chile is the country with the highest number of publicly traded stocks in the sample (88% of the listed firms) followed by Brazil (79%, but the highest absolute number of firms per country, accounting for a third of the total number of stocks), whereas Colombia posts the lowest share of listed firms covered by *Economatica* (44%). Columns 3 and 4 exhibit the mean and standard deviation of monthly-average stock returns. We observe some significant variation in mean returns and standard deviations across stock markets.

The last two columns provide estimates of the median firm market capitalization (size) and concentration.[11] In Column 5, we evidence a sizeable variability in median market capitalizations across countries. For instance, the Peruvian and Venezuelan median firms exhibit about one tenth and two tenths respectively the size of the Mexican median firm, the largest in the sample. This fact suggests that in spite of the restricted number of listed firms in each stock market, if there is any cross-sectional variation in returns associated with size we should be able to capture it. Column 6 shows a measure of market concentration, defined as the share of the equity market capitalization accumulated by the 5 largest firms. The less concentrated market is Chile (27.7%), followed by Brazil (30%).

2.3. Estimates

In this section we present and discuss the COE estimates across countries and across 21 economic sectors following the *Economatica* classification.

[10] Since major financial institutions provide widely-accepted data on emerging markets sovereign spreads in hard currency — typically in the form of bond spreads or total returns indexes, most practitioners make use of it in order to estimate modified domestic CAPM betas. However, this data has been only recently available in some Latin American countries for two reasons. First, secondary government bond markets for long maturities were nearly inexistent before the Brady Plans. Second, comprehensive indexes are only available since 1993 (e.g. JP Morgan EMBI+). Moreover, some Latin American countries like Chile had not had benchmark government bonds in hard currency until 1998. Given the constraint on sovereign spread data availability, we proceed to complete the actual dataset, namely JP Morgan EMBI + return index and spreads for Argentina, Brazil, Colombia, Mexico, Peru and Venezuela and JP Morgan EMBI Global indexes for Chile, with parametric estimations based on Druck and Morón's (2001) single equation model.

[11] We report median instead of average values because of the huge variability across firms within each stock market. For example, the ratio between the mean market capitalizations of the 10 largest and the 10 smallest firms is 536 in Argentina, 767 in Chile, and 4818 in Brazil (these figures are not reported in the table).

Table 1
Latin American stock markets—summary statistics: 1997–2004.
Source: Authors own calculations based on *Economatica*.

	Sample coverage		Stock returns		Stock valuation	
	Number of firms in the sample	Coverage rate (%)	Mean (%)	Standard deviation (%)	Median market capitalization (in millions of USD)	Market concentration (%)
Argentina	81	74	23.9	56.4	80.6	61.4
Brazil	307	79	40.0	74.0	106	30.6
Chile	211	88	15.0	32.5	103	27.7
Colombia	47	44	17.1	38.2	179	41.9
Mexico	120	51	21.2	42.1	271	39.6
Peru	115	51	17.4	32.1	34.3	48.8
Venezuela	40	74	33.8	88.2	59.8	71.3

All statistics except for those in columns 1 and 2 are computed using monthly-average observations per firm/market over the period 1997–2004. As the starting date in the sample differs across countries, e.g. it is 1986 in the case of Brazil and 1993 for Colombia, we restrict the sample to the period 1997–2004 (using information since 1993) in order to work with comparable figures. This yields 96 monthly observations per firm/country.
Columns 3 and 4 are based on a value-weighted portfolio return. The value-weighted portfolio return weighs each stock return by its respective market capitalization as a percentage of the total market capitalization. The table reports the time series means and standard deviation.

For each country and sector, Table 2 reports the volume-weighted average COE across firms.[12] Average COE estimates appear to be relatively high on an absolute basis reaching an average level of 30%.[13] However, this finding should not come as a surprise because our sample period covers several Latin American financial crisis episodes.[14]

The left panel in Fig. 1 shows that those countries posting the higher average COE estimates also display the more volatile estimates, confirming the usual positive risk-return trade-off in standard portfolio theory. This pattern reflects both the market premium that investors require in each country and the greater sensitiveness to systematic risk. Venezuela and Argentina display the higher and more volatile COE estimates, while Chile and Peru, exhibit the lowest and less volatile estimates. The Peruvian relatively lower COE and associated risk may be on account of the improved macroeconomic performance the country experienced in recent years (low and stable inflation rates and low real GDP volatility).[15]

On the right panel of Fig. 1 we show the pattern of COE estimates across sectors for Latin America as a whole. Those sectors with the highest average COE estimates are found to display the most volatile estimates.[16] Since the pattern shown in Fig. 1 is driven by the underlying beta coefficients, it illustrates to what extent systematic risk affects the COE across sectors. We find that Pension Funds and Agriculture and Fishing have the lowest cost of equity estimates in the sample, while Oil and Gas, Telecommunication, Finance and Insurance, Electricity and Construction firms carry the highest values.[17]

[12] We use traded volume (in US current dollars) instead of market capitalization as across-firm weighting variable (e.g. to obtain weighted average COE estimates by sector or country). Big firms with high market capitalization but low-trading activity (low volume) must be highly weighted when stock variables are taken into account but only high volume firms (irrespective of whether they are big or small in terms of market capitalization) should be highly weighted when flow variables are analyzed. Because COE estimates are obtained from a flow of traded shares, volume weights improve both its sector and country average reliability.

[13] It is worth noting that negative COE estimates were set to 0, what increases the latter by 25% on average. While for developed countries, negative COE estimates are rare; in emerging markets this pattern may be common during protracted crises periods where a persistent negative market premium is observed. In general, the empirical literature has dealt with negative COE estimates using a trimming rule that sets to 0 all negative COE estimates (see for example Hail and Leuz, 2004, or Barnes and Lopez, 2005).

[14] These are the Mexican crisis (1994–1995), the Brazilian crisis (1999), and the Argentine crisis of 2001–2002.

[15] Indeed, average CPI inflation in 1997–2004 stands as the lowest in Latin America and the average real GDP volatility in Peru since 1997 sits well below those of Argentina, Venezuela, Colombia, and even Mexico.

[16] Similar patterns are found when we look at each market separately.

[17] A plausible reason why Pension Funds display the lowest COE may be related with the fact that they hold more diversified investment portfolios than other sectors (with the exception of Argentina, where above 60% of their portfolio was invested in government bonds). The low COE estimates for Agriculture and Fishing may be related with the fact that this tradable sector is much more stable than other non-tradable ones. On the contrary, Construction firms are highly volatile because of their high sensitivities to the business cycle. COE estimates for Telecommunication and Electricity firms are surprisingly high because these firms have relatively stable sales (in domestic currency) and historically relatively low betas. However, expected profits in US current dollars are particularly uncertain because of country-specific regulations and high exchange rate volatility in almost Latin American countries especially from the late 90s onward.

Table 2
COE estimates, by country and sector: 1997–2004.
Source: Authors own calculations based on *Economatica*.

Sector	Argentina	Brazil	Chile	Colombia	Mexico	Peru	Venezuela	Latin America
Agriculture and fishing	0.19	0.37	0.17		0.23	0.18	0.09	0.22
	(0.26)	(0.28)	(0.17)		(0.24)	(0.21)	(0.11)	(0.24)
Chemicals and chemical products	0.24	0.28	0.09	0.21	0.22	0.15	0.25	0.27
	(0.38)	(0.34)	(0.12)	(0.21)	(0.23)	(0.29)	(0.51)	(0.35)
Construction	0.27	0.15	0.15		0.29	0.66		0.29
	(0.41)	(0.17)	(0.16)		(0.30)	(0.67)		(0.30)
Electrical equipment and electronic products	0.26	0.27	0.27			0.29		0.27
	(0.27)	(0.37)	(0.29)			(0.53)		(0.37)
Electricity	0.29	0.31	0.16	0.38		0.10	0.58	0.29
	(0.43)	(0.48)	(0.24)	(0.37)		(0.09)	1.03	(0.49)
Finance and insurance	0.36	0.24	0.13	0.27	0.31	0.16	0.25	0.28
	(0.63)	(0.35)	(0.17)	(0.37)	(0.45)	(0.25)	(0.55)	(0.44)
Food and beverages	0.29	0.21	0.22	0.35	0.26	0.22	0.15	0.25
	(0.43)	(0.26)	(0.21)	(0.47)	(0.32)	(0.38)	(0.08)	(0.32)
Machinery and equipment	0.20	0.21			0.28	0.14		0.26
	0.21	0.24			(0.34)	(0.17)		(0.31)
Mining		0.19	0.23	0.29	0.22	0.25	0.14	0.21
		(0.21)	(0.24)	(0.35)	(0.21)	(0.42)	0.11	(0.24)
Motor vehicles and related products	0.31	0.24			0.17			0.25
	(0.66)	(0.23)			(0.23)			(0.30)
Non-metallic mineral products	0.31	0.21	0.19	0.31	0.29	0.21	0.36	0.29
	0.54	(0.29)	(0.17)	(0.37)	(0.36)	(0.36)	(0.66)	(0.36)
Oil and gas	0.24	0.31	0.19	0.56				0.30
	(0.45)	(0.50)	(0.21)	(0.62)				(0.49)
Others	0.34	0.25	0.17	0.31	0.32	0.12	0.19	0.29
	(0.47)	(0.36)	(0.21)	(0.34)	(0.38)	(0.21)	(0.37)	(0.38)
Paper and paper products	0.22	0.20	0.21	0.10	0.26		0.30	0.23
	(0.20)	(0.27)	(0.24)	(0.07)	0.29		0.56	(0.28)
Pension funds			0.18			0.10		0.17
			(0.23)			(0.01)		(0.22)
Software and data	0.22			0.12				0.21
	(0.25)			(0.04)				(0.24)
Steel and metal products	0.41	0.29	0.42		0.29	0.37	0.52	0.30
	(0.90)	(0.38)	(0.43)		0.34	0.52	1.01	(0.44)
Telecommunications	0.40	0.31	0.27	0.04	0.20	0.16	0.29	0.29
	(0.73)	(0.40)	(0.35)	(0.07)	(0.29)	(0.29)	(0.64)	(0.46)
Textiles	0.20	0.21	0.07	0.26	0.06	0.14	0.52	0.22
	(0.24)	(0.24)	(0.07)	(0.27)	(0.12)	(0.18)	(0.99)	(0.32)
Transportation and storage	0.05	0.12	0.25		0.23		0.17	0.22
	(0.00)	(0.11)	(0.23)		(0.16)		(0.34)	(0.21)
Wholesale and retail trade	0.18	0.33	0.34	0.18	0.24	0.22		0.26
	(0.29)	(0.43)	(0.31)	(0.16)	(0.30)	0.36		(0.32)
All sectors	0.32	0.29	0.21	0.33	0.18	0.19	0.47	0.3
	(0.58)	(0.45)	(0.25)	(0.37)	(0.32)	(0.30)	(0.79)	(0.47)

For each firm in the sample we estimate the COE based upon the adjusted version of CAPM stated in Eq. (3) in Section 1. For each country and sector, the first row reports the volume-weighted average COE estimation. The second row displays its standard deviation (in parentheses).

2.4. Accounting for the sources of risk

How much of the average stock excess return (COE minus the risk-free rate) can be attributed to systematic risk in Latin American markets? Our econometric estimations find that on average a 33% of the variability in the typical Latin American stock excess return is attributable to systematic risk.

We get this result devising two alternative measures of the explained variance in the GMM regressions of our adjusted CAPM equation (Eq. (2)). The first measure is the regression adjusted-R^2 or goodness of fit which isolates the percentage of the total risk which is accounted for by systematic risk (as opposed to the

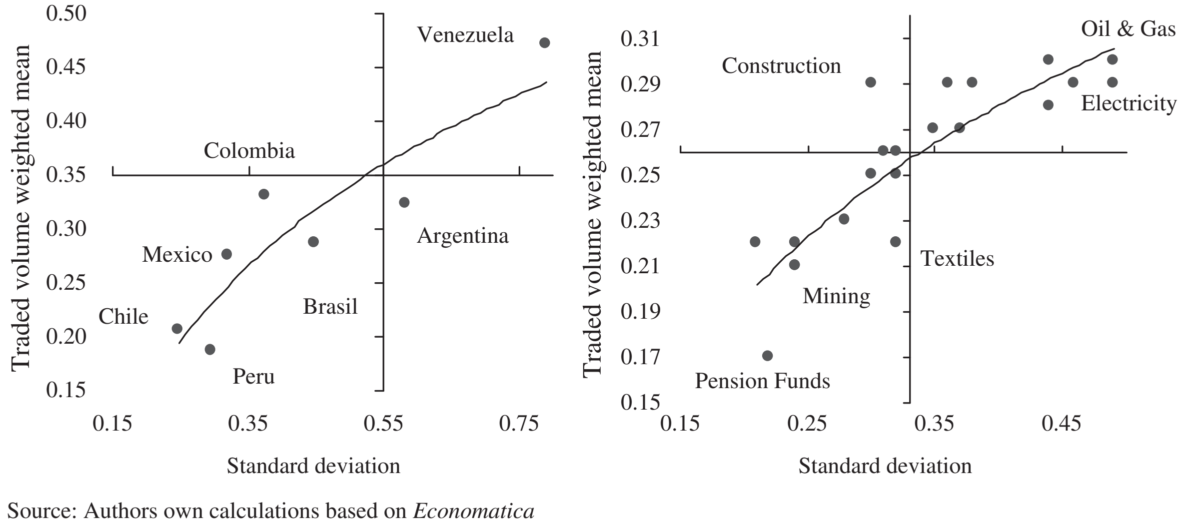

Fig. 1. Cost of equity estimates by country and sector (1997–2004).
Source: Authors own calculations based on *Economatica*.

residual, i.e. unique risk). As stated earlier, the CAPM assumption on the efficiency of the market portfolio implies that it should fully capture the systematic component of risk. However, there are some disadvantages to using this measure, e.g. it is model dependent and is only a relative measure.

A caveat is in order: the variance of COE equals the variance of the stock excess return adjusted by sovereign risk, in turn equal to systematic risk or the variance of stock market excess returns adjusted by sovereign risk too only if the variance of the risk-free rate, which in our exercise is proxied by the variance of a UST bond yield, is zero, and beta is a constant at period t.[18] Remember all excess returns are expressed in US current dollars.

Goyal and Santa Clara (2003) provide an alternative variance decomposition procedure. Basically the methodology consist in first estimating the cross-sectional variance of stock (excess) returns in a given country at a certain moment in time to capture the systematic component of stock (excess) return volatility, and, second, to divide the latter by the average total stock risk.[19]

The results for both measures are presented in Table 3. They are notably similar suggesting that, for Latin America as a whole, systematic risk accounts for nearly a third of stocks excess returns (and COE) variability. Argentina (43%, on average) and Venezuela (39%) stand with the greater share of systematic risk, consistently across measures. The relatively lowest share of systematic risk is found in Peru (25%).

[18] Formally:

$$VAR\left[E\left(R_{i,t}-R_{f,t}-SS_{j,t}\right)\right] = VAR(COE_{i,t}) = \beta_i^2 VAR\left[R_{M,t}-R_{f,t}-SS_{j,t}\right]$$

because $VAR(R_{f,t})=0$ and β_i^2 is constant at time t (and therefore $COV(R_{f,t},R_{M,t}-R_{f,t}-SS_{j,t})=0$).

[19] Consider the following measure of total risk in a country. First, compute the variance of a stock (or collection of stocks) p using a 12 months moving window as:

$$V_{pt} = \sum_{i=0}^{11} r_{pt-i}^2$$

where $r_{p,t-i}$ is stock p return in month t. Then compute the average stock variance as the arithmetic average of the 12-month window variances of each stock return.

$$V_t = \frac{1}{N_t}\sum_{i=1}^{N_t}\left[\sum_{i=0}^{11} r_{pt-i}^2\right]$$

where N_t is the number of stocks in the country of interest at each t. Analyzing together both measures, we derive the share of the total variance of returns explained by idiosyncratic risk, namely $1-\frac{V_{ew}}{V_t}$. With the size of this "residual" variable (unique risk) we obtain a measure of the potential mis-pricing error in CAPM models if at least one of the underlying assumptions does not hold.

Table 3
Measures of the share of systematic risk by country (1997–2004).
Source: Authors own calculations based on *Economatica*.

Country	Measure 1: goodness of fit	Measure 2: variance decomposition (Goyal and Santa Clara)	Simple average
Argentina	44%	41%	43%
Brazil	23%	33%	28%
Chile	32%	28%	30%
Colombia	40%	25%	33%
Mexico	44%	30%	27%
Peru	29%	21%	25%
Venezuela	41%	37%	39%
Average	36%	31%	33%

Measure 1: The goodness of fit is the average adjusted R-squared for each N stocks in country *i*.
Measure 2: The systematic risk share in total risk is calculated dividing the cross-section variance of stock *j* over the relevant period by the average measure of total risk over the same period (see Goyal and Santa Clara, 2003).

In a cross-country comparison, all these shares remain well above those obtained from many North American, European and Asian stock markets. In the case of the US, Goyal and Santa Clara (2003) and Fu (2009) show that idiosyncratic risk accounts for more than 80% of total risk (excess return total variance). Using a broader database, Guo and Savickas (2007) obtained similar results for United Kingdom, Canada and Japan.

Notice that, under the CAPM assumptions, the remaining two-thirds of the unexplained variance should be attributable to idiosyncratic, fully diversifiable risk. However, investors might not be able to fully diversify portfolios (as Levy, 1978, among others, demonstrates). Diversification opportunities may be scarce in relatively shallow and illiquid markets including for Latin America as recent evidence points out.[20] As a consequence, the part of the excess return and COE variance that remains unexplained could reflect a potential mispricing error in the cost of equity capital.

3. Robustness checks

3.1. Test for international risk factors

In the decade to 2005 Latin American stock markets exhibited an increasing degree of co-movement and correlation with the most representative global markets. This phenomenon raises the question of whether Latin American markets are increasingly globally integrated, and in particular, to what extent a change in the assumptions regarding the integration of markets changes the estimates of the COE.

We test for the inclusion of global factors[21] into the adjusted CAPM equation estimated in Sections (2–3) following Koedijk et al. (2002). They set forth a multifactor stock pricing model close to Eq. (7) below:

$$R_{i,t} = \alpha_i + \beta_{iM}\left[R_{M,t}\right] + \beta_{iG}\left[R_{G,t}\right] + \beta_{i\mathrm{MNER}}\left[R_{\mathrm{MNER},t}\right] + \varepsilon_{i,t} \tag{7}$$

where $\left[R_{G,t}\right]$ is the global market portfolio return, $\left[R_{\mathrm{MNER},t}\right]$ is the nominal multilateral exchange rate return, and β_{iG} and $\beta_{i\mathrm{MNER}}$ are the associated asset *i* return sensitivities or global factor loadings. The key idea is that there might be additional sources of systematic risk in the pricing of stocks not captured by local portfolios, if the firm's stock excess returns are correlated with excess returns on internationally diversified portfolios of stocks or currencies.

The null hypothesis to be tested is $H_0: \beta_{i\mathrm{MNER}} = \beta_{iG} = 0$. A rejection of H_0 would support a change in the specification of the domestic adjusted CAPM model so as to incorporate global risk factors into Eq. (3). Notice

[20] See de la Torre et al. (2008).
[21] Solnik (1974), Sercu (1980, 1981) and Adler and Dumas (1983) demonstrated that in a fully mobile capital world investors hold internationally diversified portfolios of risky assets and regard the risky security choices by how (in terms of risk and returns) they contribute to their internationally diversified portfolios.

Table 4
Percentage of tests that reject the null hypothesis of non significant global factors.
Source: Authors own calculations based on *Economatica*.

Country	Global portfolio return	Multilateral real exchange rate return	Joint test	Period
Argentina	6	15	16	1990–2004
Brazil	14	23	28	1986–2004
Chile	5	15	14	1989–2004
Colombia	7	8	11	1993–2004
Mexico	6	23	18	1991–2004
Peru	7	7	8	1992–2004
Venezuela	8	10	14	1989–2004
Latin America	9	18	20	1986–2004

The rejection rates are calculated on the basis of volume-weighted GMM models using the number of monthly traded shares as within weights.

that all variables are expressed in absolute and not in excess returns over a risk-free rate, i.e. we rescale all variables. Finally, we prefer to use multilateral real exchange rate returns instead of using nominal exchange rate returns in order to allow for deviations from the Purchasing Power Parity (PPP) rule and a partial correlation of local returns with inflation rates. As a result, we estimate Eq. (8) below for each of the 921 firms in our sample over each of the (rolling) periods starting from the earliest observation possible:

$$R_{i,t} = \alpha_i + \beta_{iM}\left[R_{M,t}\right] + \beta_{iG}\left[R_{G,t}\right] + \beta_{i\mathrm{MRER}}\left[R_{\mathrm{MRER},t}\right] + \varepsilon_{i,t} \tag{8}$$

where $R_{\mathrm{MRER},t}$ stands for the return on a multilateral real exchange rate portfolio. Now $H_0 : \beta_{i\mathrm{MRER}} = \beta_{iG} = 0$

We report the percentage of cases in which we reject H_0 in Table 4 below. Rejection rates are small, averaging between 20% and 29% depending on the specification.[22] This implies that the global market portfolio and multilateral real exchange rates are not jointly statistically significant variables and should be disregarded as potential additional systematic risk factors in the determination of Latin America's COE.

We conclude that global factors do not add explanatory power to domestic adjusted CAPM regressions. Clearly, this result does not suggest that Latin American markets are isolated. In light of the increasing pattern of co-movement that the region stock markets exhibit, our tests instead might indicate that a great deal of the country-specific or regional risk is indeed correlated with global risk and, at least over the time span we cover, the local market portfolio is sufficient to capture global sources of non-diversifiable risks affecting local stocks.

3.2. Testing the Fama and French three factor model: Size and value premia

To check for the omission of other sources of systematic risk as put forth by Fama and French (1992, 1993, 1996, 1998) and its applicability to our sample of Latin American stocks we proceed in the following way: First, we search for the presence and persistence of size (small firms carry higher expected returns than big firms) and value premia (distressed firms pay off higher returns than growth firms) in those stocks and compare the results with those found by Fama and French (1998) and Rouwenhorst (1999). Then, we rerun the basic adjusted CAPM adding the premia on size and value sorted portfolios to the local market risk premium.

3.2.1. Descriptive statistics

To compute the size and value premia we follow the procedure adopted by Fama and French (1998). For each year in our sample, we sort stocks according to their mean market equity capitalization or to their book-to-market value and group them into three portfolios: the top 33.3%, middle 33.3% and bottom 33.3%. Then, we drop the medium portfolios, so we only keep the big size (B), small-size (S), high book-to-market (H) and low book-to-market (L) portfolios. Finally, we calculate the monthly value-weighed return on each portfolio (B, S, H and L respectively).

[22] Although we do not report the results in this paper, our tests exhibit increasing rejection percentages since 2000 and average p-values slightly decreasing over time.

Table 5
Annual dollar returns in Latin American stock markets, size and book-to-market ratios sorted portfolios: 1997–2004.
Source: Authors own calculations based on *Economatica*.

Country	Domestic market (I)	Book-to-market (value)			Market capitalization (size)		
		H (II)	L (III)	H–L (IV)	Small (V)	Big (VI)	S–B (VII)
Argentina	0.24	0.24	0.23	0.01	0.13	0.24	−0.11
	(0.01)	(0.13)	(0.09)	(0.04)	(0.00)	(0.13)	(−0.08)
	[0.56]	[0.58]	[0.64]	{0.17}	[0.48]	[0.57]	{−1.39}
Brazil	0.40	0.38	0.41	−0.04	0.43	0.40	0.03
	(0.03)	(0.27)	(0.19)	(0.06)	(0.36)	(0.29)	(0.09)
	[0.74]	[0.71]	[1.05]	{−0.28}	[0.55]	[0.75]	{0.31}
Chile	0.15	0.15	0.16	−0.01	0.09	0.15	−0.06
	(0.01)	(0.11)	(0.06)	(0.03)	(0.06)	(0.09)	(0.01)
	[0.33]	[0.32]	[0.42]	{−0.24}	[0.20]	[0.33]	{−1.53}
Colombia	0.17	0.17	0.17	0.01	0.21	0.17	0.03
	(0.02)	(0.14)	(0.11)	(0.02)	(0.16)	(0.15)	(0.01)
	[0.38]	[0.37]	[0.43]	{0.12}	[0.39]	[0.40]	{0.50}
Mexico	0.21	0.22	0.19	0.02	0.14	0.21	−0.08
	(0.02)	(0.13)	(0.14)	(0.07)	(0.10)	(0.14)	(−0.05)
	[0.42]	[0.44]	[0.40]	{0.41}	[0.30]	[0.44]	{−1.40}
Peru	0.17	0.18	0.30	−0.12	0.26	0.17	0.09
	(0.00)	(0.15)	(0.17)	(−0.05)	(0.22)	(0.13)	(0.08)
	[0.32]	[0.35]	[0.55]	{−1.82}	[0.38]	[0.33]	{1.78}
Venezuela	0.34	0.31	0.22	0.10	0.25	0.34	−0.10
	(0.00)	(0.00)	(0.02)	(0.06)	(0.05)	(0.04)	(0.03)
	[0.88]	[0.93]	[0.79]	{0.71}	[0.59]	[0.93]	{−0.73}

Column I displays USD annual average returns on local market portfolios. Columns II through VII report USD annual average return on portfolios sorted by book-to-market ratios and market capitalization. For each country, the first row reports the value-weighed portfolio return. The second row displays the median return (in parentheses). The third row reports the standard deviation (in brackets) or the *t*-statistic of the mean difference tests (in braces, Columns V through VII).

Table 5 reports the mean, median and standard deviation of each of the portfolios returns over 1997–2004. For benchmarking purposes, we also include the local market value-weighed portfolio in the first column. If small stocks (S) persistently outperformed big stocks (B) and value stocks (H) persistently outperformed growth stocks (L), then the mean (or median) return differences $S-B$ and $H-L$ should be positive and statistically different from zero. Column 4 in Table 5 ($H-L$) shows positive $H-L$ mean and median return differences in only 4 of the 7 countries. However, none of them are statistically significantly different from zero. Furthermore, in the sole case where the $H-L$ mean return is significantly different from zero (Peru: t-value $= -1.82$), $H-L$ turns out to be negative, contrary to our expectation. Similarly, we only find positive $S-B$ mean and median returns in two countries, namely Brazil and Peru but only the latter posts an $S-B$ mean return significantly different from zero (t-value: 1.78). In conclusion, there is no robust evidence of a presence and persistence of size and value premia in Latin American stocks over 1997–2004.

Table 6 exhibits the value ($H-L$) and size ($S-B$) premia found in Fama and French (1998), Rouwenhorst (1999) and the present study altogether. We only report figures for the stock markets sampled in our study. Rouwenhorst (1999) finds positive value premia in 5 of 6 countries, yet only one comes out statistically significant (Brazil). Fama and French (1998) find 3 stock markets displaying positive value premia but none is significantly different from zero. As to the size premia, the pattern is similar. Rouwenhorst (1999) reports positive size premia in 5 out of 6 countries yet only two among them pass the t-test (Argentina and Mexico). Fama and French (1998) also find positive size premia for the same markets (with the exception of Colombia), but none is statistically significant.

A caveat is in order. We often observe that either the $H-L$ or the $S-B$ "excess" returns yield opposite signs across studies or are statistically significantly different from zero in one study while not in the others. There are at least two limitations to this comparative analysis: 1) the sample period and the data frequency[23] do not

[23] Our results do not change when we allow for the same data frequency as in Fama and French (1998) or Rouwenhorst (1999). These figures are available from the authors upon request.

Table 6
Size and value premia across three studies.
Source: Authors own calculations based on *Economatica*.

Country	Book-to-market			Market capitalization		
	H–L			*S–B*		
	Rouwenhorst, 1999	Fama and French, 1998	This paper	Rouwenhorst, 1999	Fama and French, 1998	This paper
	1982–1997	1987–1995	1997–2004	1982–1997	1987–1995	1997–2004
Argentina	1.68	−0.36	0.01	3.84*	0.04	−0.11
Brazil	3.94*	0.73	−0.04	1.76	0.12	0.03
Chile	1.07	0.15	−0.01	0.31	0.09	−0.06
Colombia	−0.36	−0.17	0.01	−0.68	−0.21	0.03
Mexico	1.39	0.00	0.02	2.39*	0.03	−0.08
Peru	Na	Na	−0.12*	Na	Na	0.09*
Venezuela	1.27	0.57	0.10	1.85	0.24	−0.10

This table replicates the evidence on value and size premia for 7 Latin American stock markets from Fama and French (1998) and Rouwenhorst (1999) and contrasts their results to ours. *H–L* and *S–B* portfolios are formed as explained in Section 3.2.1 with some minor differences about the frequency of the sorting of portfolios: Fama and French (1998) sort stocks using information at the end of each year, Rouwenhorst (1999) uses monthly information, and this paper uses annual averages. An * (asterisk) indicates statistically significant differences between mean returns.

coincide; and 2) the number of firms covered in each study is not the same. For instance, it has been documented in Latin America that a great deal of firms delisted more remarkably during the second half of the 1990s (see de la Torre et al., 2008). Notwithstanding this, all the three studies point to the lack of generalized robust evidence of presence and persistence of size and value premia in Latin American stocks throughout the period 1982–2004.

3.2.2. Econometric regression results

The question now is whether size and value risk premia are statistically significant variables in our adjusted CAPM model, even if for most Latin American countries the mean difference between big and small or high and low book-to-market value stock returns are not statistically different from zero as demonstrated in Section 3.2.1.

As said before, Fama and French (1993) introduce an extended CAPM model (FF3FM, Eq. (9) below), where a firm's excess return over the risk-free rate is explained by the excess return on the market portfolio (CAPM systematic risk) and two additional factors designed to capture other sources of systematic risk apart from market portfolio risk: SMB (small minus big), which is the difference between the returns on well-diversified portfolios of small and big stocks or simply "size premium", and HML (high minus low), which is the difference between the returns on well-diversified portfolios of high and low book-to-market stocks or "value premium". More specifically:

$$E\left(R_{i,t}\right)-R_{f,t} = \beta_{iM}\left[E\left(R_{M,t}\right)-R_{f,t}\right] + \beta_{iS}[E(SMB_t)] + \beta_{iH}[E(HML_t)] \tag{9}$$

with

$$SMB_t = \sum_{j=\underline{\underline{j}}}^{\underline{j}} R_{j,t}w^S_{j,t} - \sum_{j=\overline{j}}^{\overline{\overline{j}}} R_{j,t}w^B_{j,t} \tag{10}$$

where $j\in\left[\underline{\underline{j}}, \overline{\overline{j}}\right]$ is a firm index, $\underline{\underline{j}}$ is the smallest firm (the company posting the lowest market capitalization), $\overline{\overline{j}}$ is the largest one, $\underline{j}$ is the upper bound firm of the small-size group (i.e. the 30th percentile firm), $\overline{j}$ is the lower bound firm of the big size group (the 70th percentile firm), $w^S_{j,t} = MKC_{j,t} / \sum_{j=\underline{\underline{j}}}^{\underline{j}} MKC_{j,t}$ and $w^B_{j,t} =$

$MKC_{j,t} / \sum_{j=\bar{j}}^{\bar{\bar{j}}} MKC_{j,t}$ are the market capitalization shares of firms $\left[\underline{\underline{j}}, \underline{j}\right]$ and $\left[\bar{j}, \bar{\bar{j}}\right]$, respectively (with $MKC_{j,t}$ being the firm j market capitalization at t); and

$$HML_t = \sum_{k=\bar{k}}^{\bar{\bar{k}}} R_{k,t} w_{k,t}^{H} - \sum_{k=\underline{\underline{k}}}^{\underline{k}} R_{k,t} w_{k,t}^{L} \tag{11}$$

where $k \in \left[\underline{\underline{k}}, \bar{\bar{k}}\right]$ is another firm index, $\underline{\underline{k}}$ is the firm displaying the lowest book-to-market ratio, $\bar{\bar{k}}$ is the firm with the highest book-to-market ratio, $\underline{k}$ is the 30th percentile firm, $\bar{k}$ is the 70th percentile firm, while $w_{k,t}^{H}$ and $w_{k,t}^{L}$ are defined in the same way as $w_{j,t}^{B}$ and $w_{j,t}^{S}$.

For each firm/country, we test the null hypothesis $H_0: \beta_{iS} = \beta_{iH} = 0$ performing standard Wald tests on GMM estimations of β_{iS} and β_{iH} from Eq. (9). A rejection of H_0 would imply that either SMB or HML or both are significant systematic risk factors and therefore should not be omitted in the pricing of firm's stocks and subsequent calculation of the COE. If that was the case, we would expect $\beta_{iS} > 0$ and $\beta_{iH} > 0$. By contrast, the non-rejection of H_0 would lead us to conclude that either SMB or HML or both do not add information to the domestic market portfolio in the explanation of firm's stock excess returns.

Table 7 summarizes the percentage of Wald tests per country where the null hypothesis of insignificant FF factors is rejected at the 5% level. Overall, we conclude that the FF factors do not carry significant information for stock pricing in Latin America. On average, the percentage of Wald test rejection reaches 40% in Brazil (i.e. in 40% of the regressions we reject the null $H_0: \beta_{iS} = \beta_{iH} = 0$) whereas it is generally below 30% in the other Latin American countries. The relatively higher proportion of rejections in Brazil does not come as a surprise as 1) we observe the highest standard deviation of firm's size in Brazil (relative to other Latin American countries), which renders size a more meaningful risk factor; 2) there is a relatively low explanatory power of CAPM in this country (see Section 2); and 3) value premia turns out significant in the FF3FM model run by Rouwenhorst (1999) over Brazilian stocks. Table 8 presents the dynamic properties of the Wald test results, where no clear-cut time trend is observed in most Latin American stock markets. Nevertheless, it must be noticed that the rejection rates of the null hypothesis (non significant FF factors) are always below the 50% excluding Brazil in 2003.

While the calculation of SMB and HML and their statistical significance through a mean difference test are comparable with those of Fama and French (1998) and Rouwenhorst (1999) despite the stock coverage and time period not being equal, the econometric methodology and results differ in at least three respects: 1) Fama and French (1998) estimate the cross-section of expected returns including emerging and developed economies whereas we estimate time series pricing equations, one per stock/firm in Latin America; and 2) Rouwenhorst (1999) captures both the cross-section and time series dimension of a sample comprising 20 emerging markets including ours but her estimator is not a rolling-window GMM and her stock returns are not adjusted for illiquidity, interest rate volatility and sovereign risk in the way we follow.

Table 7
Proportion of Wald tests rejecting the null hypothesis of non significant Fama–French factors.
Source: Authors own calculations based on *Economatica*.

Country	*S–B* (size premia)	*H–L* (value premia)	Joint test	Period
Argentina	0.12	0.12	0.18	1990–2004
Brazil	0.29	0.25	0.40	1986–2004
Chile	0.15	0.16	0.25	1989–2004
Colombia	0.15	0.20	0.31	1993–2004
Mexico	0.13	0.16	0.23	1991–2004
Peru	0.16	0.12	0.22	1992–2004
Venezuela	0.15	0.15	0.25	1989–2004
Latin America	0.20	0.19	0.31	1986–2004

Rejection rates are computed on the basis of weighted GMM regressions using monthly traded shares as within weights.

Table 8
Proportion of Wald tests rejecting the null hypothesis of non significant joint Fama–French factors.
Source: Authors own calculations based on *Economatica*.

Year	AR	BR	CL	CO	MX	PE	VE	Lat. Am.
1990		0.31						0.31
1991		0.35						0.35
1992		0.45						0.45
1993		0.38	0.19					0.31
1994		0.39	0.21				0.29	0.32
1995		0.36	0.19		0.38		0.04	0.29
1996	0.20	0.28	0.16		0.22	0.35	0.24	0.24
1997	0.19	0.32	0.17	0.21	0.14	0.29	0.22	0.25
1998	0.20	0.32	0.18	0.39	0.13	0.17	0.21	0.23
1999	0.16	0.40	0.29	0.49	0.19	0.12	0.23	0.29
2000	0.18	0.42	0.32	0.35	0.22	0.15	0.17	0.30
2001	0.18	0.43	0.35	0.24	0.22	0.20	0.25	0.32
2002	0.15	0.47	0.32	0.25	0.25	0.28	0.33	0.34
2003	0.12	0.55	0.30	0.34	0.28	0.26	0.31	0.36
2004	0.23	0.48	0.26	0.25	0.31	0.21	0.33	0.34
Full Simple average	0.18	0.40	0.25	0.31	0.23	0.22	0.25	0.31

Yearly average by country obtained from weighted GMM regressions, using monthly traded shares as within weights.

In general, our findings are in line with other recent studies e.g. Liu (2006) who, controlling for liquidity risk, find considerable evidence of the limited explanatory power of the Fama and French model to capture the cross-section of expected stock returns in the NYSE, AMEX and NASDAQ markets. Martinez et al. (2005) also present evidence of the limited explanatory power of the Fama and French three factor model using the Spanish stock market as a case study, although they show there is some evidence of some explanatory power in retaining the size factor.

4. Summary and conclusions

This paper aimed to answer the following questions: which are the sources of risk driving Latin American stock prices and its corresponding opportunity cost in the period 1993–2004? How can the best fit of the opportunity cost of equity be obtained within the standard asset pricing framework?

Taking as starting point an adjusted version of CAPM, we first investigated how much of COE variability could be attributed to systematic (i.e., non-diversifiable) risk. We found that for Latin America as a whole, nearly a third of COE variability on average can be attributed to systematic sources of risk (between 10 and 78% more than in stock markets such as the US, UK, Canada and Japan, (see Fu, 2009; Guo and Savickas, 2007).

In a second step, we checked the robustness of these results to the inclusion of global stock and real currency portfolios. Here the question was: should internationally well-diversified portfolios be factored in the pricing of Latin American stocks? In fact, as Latin American countries opened up its capital markets to foreign investors and let its residents invest abroad, the question of how much of the individual stock risk borne by Latin American firms is correlated with global risk factors became pervasive (see Bekaert and Harvey, 2000; or Henry, 2000). Our results, obtained using data corresponding to the post-equity market liberalization period in Latin America suggest that global factors do not add explanatory power to domestic portfolios in the adjusted CAPM regressions, which may signal that a significant fraction of regional/country-specific stock market risk is indeed correlated with global risk and hence that local market portfolios are already capturing the relevant sources of global risk. They also indicate that increasing but still modest correlations between developed country and Latin American stock excess returns (except for Mexico, whose correlation with US stocks is somewhat higher) may reduce the informational power of global portfolios and the diversification benefits of the latter from the domestic investor perspective (Bekaert and Harvey, 2003). Or thinking the other way round, should we have tested an international adjusted CAPM (Sercu, 1980; Solnik, 1974) and then added the domestic market in the pricing equation the latter would have come out statistically and economically redundant. Other potential explanations for these findings could be the presence of a home bias or country barriers to investment in foreign stocks such as taxes, capital controls (Chile until 2000 or Argentina after the 2002 default and devaluation) and other market distortions that would prevent local investors to diversify their portfolios internationally.

In a third and last step, we tested extended our analysis over the same sample stocks where global risk factors had come out insignificant in our adjusted CAPM GMM regressions to test for the omission of other sources of common risk originated in the size and book-to market value of stocks that pervade in their returns and that domestic (and global) market portfolios fail to price in. This is the Fama–French Three Factor Model (FF3FM, Fama and French, 1992, 1993, 1996, 1998). We concluded that: 1) both size and value premia are not generally statistically significant risk factors, and 2) they do not add informational power to the domestic market portfolio in the explanation of stock (excess) returns, with the exception of Brazil for some years.[24] These findings are in line with Fama and French (1998), Liu (2006) and Martinez et al. (2005) and to some extent with Rouwenhorst (1999) because the latter found a global significant effect of both premia on a cross-section of emerging market excess returns.

The limited or no explanatory power of the FF3FM variables is less controversial than the role of global risk factors and other conditioning variables potentially play in the pricing of emerging market stocks like those of Latin America. As stated by Barclay et al. (2010) and mentioned above, Bekaert and Harvey (2000, 2002) find an increase in correlation between emerging country (EM) and developed capital markets in the post-liberalization period. They show that EM betas with the world portfolio increase on average two and a half times between the pre- and post-liberalization periods (focus of this paper), indicating significant increased EM responsiveness to global market risk. Despite these ostensibly substantial increases, Bekaert and Harvey (2002) nevertheless conclude that EM correlations with the developed world are still low enough to the global investor with significant portfolio diversification. However, other authors like Fernandez (2003), cited in Barclay et al. (2010) suggest that EM stocks' exposure to the global market is no longer significantly different from that of developed market stocks.

Furthermore, several recent studies have shown a number of other important factors in EM stock market returns including time variation, local information variables and currency risk. For instance, Carrieri et al. (2007) examine the relevance of both time-varying global and local market risk on the expected returns and find evidence of significant time variation in the price of both local and global market risk. This suggests that specifying a global asset pricing model that holds period-by-period significantly improves the mean-variance efficiency of the world portfolio, and thus the model's ability to explain EM returns. These findings go counter to ours but may have to do with their sample period, estimation method and the adding of local conditional variables which are absent in our econometric exercise. Finally, in terms of currency risk Carrieri et al. (2006) find that EM currency risk (measured in real terms) is priced separately from other EM risks and represents a significant proportion of equity returns in both developed and EM, again contrary to our predictions for Latin American stocks.

In short, an adjusted version of CAPM which takes into consideration the lack of liquidity in Latin American stock markets, the instability of betas over time, the volatility in interest rates in an environment where these rates are not strictly risk-free and the across-sector dispersion of those betas turns out a “second” best model to estimate the firm's cost of equity. Second best because there may be other sources of systematic risk, including for example systematic liquidity (see Hearn and Piesse, 2009) which alternatively we consider implicit in or specific to each stock, contagion risks, or industry-specific variables (Carrieri et al., 2007; Barclay et al., 2010) that could improve the fit of our GMM COE regressions. Recall we estimate the time series adjusted CAPM COE for 921 stocks and we do not perform at any stage the cross-section estimations of the various betas, i.e. domestic, global, size and book-to-market, nor do we compute their corresponding fitted excess returns.

Finally, we should caution against the use of these COE estimates to value common stock of Latin American firms, let alone their fair market value (see Bruner et al., 2002, for a discussion of alternative valuation methods for stocks, including for Latin American firms or Pereiro, 2006, for the specific case of Argentina). Further work may consider other conditional versions of CAPM (Bonomo, 200; Hearn and Piesse, 2009; Iqbal et al., 2010; Barclay et al., 2010) or downside beta/semi-variance portfolio models (Galagedera, 2007) to estimate the cost of equity and assess their explanatory power compared to our regression results.

[24] Further regressions not reported in this paper including global and FF3F model risk factors altogether yield the same conclusion, that is, in general domestic portfolio risk suffices to explain the time variation of stock excess returns. These results are available from the authors upon request.

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