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The anatomy of financial crises: Evidence from the emerging ADR market ☆

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ABSTRACT

We study the anatomy of recent financial crises in Mexico, East Asia, Russia, Brazil, Turkey, and Argentina by investigating the efficiency and pricing of the emerging American depositary receipt (ADR) market. We use a non-parametric technique to test for persistent regime shifts in two basic structural relationships for ADR returns in 20 emerging countries — identified via arbitrage and capital mobility considerations — that should always hold in efficient and integrated capital markets. We find that those "normal" market conditions were instead often violated in proximity of financial crises: The law of one price often weakened (by 54% on average) and domestic sources of risk became more important (often by more than 100%) for many emerging ADRs. We also find the likelihood of these regime shifts to be related to proxies for uncertainty among investors, exchange rate volatility, trade linkages, and liquidity (but not stock market trends, currency devaluations, capital flight, or capital controls).

1. Introduction

The global financial markets have recently been plagued by recurrent episodes of turmoil, most notably originating in Mexico at the end of 1994, in East Asia between 1997 and 1998, in Russia during the summer of 1998, in Brazil in 1999, in Turkey in 2001, and in Argentina in 2001/2002. In proximity of their occurrence, most emerging and some developed markets experienced sudden, severe, often deemed "excessive," downward movements in stock, currency, and sovereign debt prices, scarce or drying liquidity, rapid reversals of capital flows, significant output losses, bank runs, or spillover effects.

The concentration of these episodes of turbulence, usually referred to as financial crises, over a relatively short period of time has drawn the attention of a large and growing theoretical and empirical literature. Financial crises have alternatively been attributed to local macroeconomic and microeconomic weaknesses (e.g., Krugman, 1979; Agenor et al., 1992; Kaminsky et al., 1998), monetary policy (e.g., Aghion et al., 2001), coordination problems among investors (e.g., Chang and Velasco, 1999), the activity of large traders and speculators (e.g., Brown et al., 2000; Kaminsky et al., 2001, 2004; Kyle and Xiong, 2001; Kim and Wei, 2002; Kodres and Pritsker, 2002; Corsetti et al., 2004), herding (e.g., Chari and Kehoe, 2004), self-fulfilling equilibria (e.g., Obstfeld, 1998; Flood and Marion, 2000), the interaction of stock and foreign exchange markets (Corsetti et al., 1999; Park and Lee, 2003), or contagion (e.g., King and Wadhwani, 1990; Calvo and Mendoza, 2000a,b; Corsetti et al., 2005; Kallberg et al., 2005; Pasquariello, 2007).[1]

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[1] For economy of space, the above list is less than comprehensive. Calomiris (1995), Buiter et al. (1998), Corsetti et al. (1999), Flood and Marion (1999), Jeanne (2000), and Kaminsky et al. (2003) provide extensive surveys of this literature and discuss its most recent developments.

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Operational criteria for the identification of the main, common characteristics of financial crises are a necessary premise of these studies (e.g., Mishkin, 1992). According to Bordo et al. (2000, p. 4), "any study of financial crises will turn on how these events are identified." Thus, for instance, defining crises as episodes of financial turmoil followed by output declines may lead to ignore instances when such turmoil did not have real effects. Alternatively, Schwartz (1986) observes that many financial disturbances leading to falling asset prices should not be viewed as financial crises — and labeled instead "pseudo-financial crises" — if they are limited to a particular sector of the economy or set of market participants. In addition, identifying the nature of crisis events is arguably a crucial step in policymakers' and regulators' efforts at formulating prescriptions to prevent their occurrence or attenuate the severity of their effects.

Typically, financial crises are identified — i.e., distinguished from tranquil periods — and dated using various statistical criteria based, e.g., on large fluctuations in stock and real estate prices, foreign exchange market pressure, a protracted deterioration in banks' loan portfolios, bank failures, or common illiquidity and insolvency problems among financial market participants (Eichengreen et al., 1996; IMF, 1998; Bordo et al., 2000). The objective of this paper is to provide further evidence on the anatomy of recent financial crises by assessing the impact of their occurrence on the efficiency and integration of emerging equity markets. Besides their distinct importance (e.g., Karolyi, 1998, 2006; Bekaert and Harvey, 2003), both features are likely to interact in the context of those markets, e.g., by virtue of the activity of foreign investors. For our purpose, we focus on the market for emerging American depositary receipts (ADRs). ADRs are dollar-denominated rights, priced in the U.S. stock market, to shares of foreign companies traded in their domestic exchanges. According to Bailey et al. (2002), ADRs account for much trading in emerging market equities. Depositary receipts allow investors to trade in more transparent and liquid markets (the NYSE, NASDAQ, or AMEX) than those of the issuers. More important, the ADR market represents an ideal environment to determine whether efficiency and integration of emerging equity markets deteriorates during periods of financial stress. In each of those markets, ADRs and their corresponding amounts of shares of the domestic issuers should always be perfect substitutes. Most of the available research on the subject (see Section 2) argues that this is generally the case, in the absence of investment restrictions, when taxes, transaction costs, liquidity, and non-synchronicity issues are considered. This implies that the relationship between the dollar return for a depositary receipt and the dollar return for the underlying security should always satisfy the law of one price, even after accounting for those financial frictions, unless "normal" market conditions are violated.

Further, in perfectly open, fully integrated financial markets, ADR prices should be related exclusively to their covariance with world factors. Most emerging economies have experienced some form of capital market liberalization in the last decade. As a result of this process, foreign investors can now trade in domestic assets and domestic investors in foreign assets. However, it is well known that investors' portfolios still display significant home bias (e.g., Tesar and Werner, 1995a); hence, local risk factors may be reasonably expected to affect ADR prices to a large extent. Importantly for our analysis, these factors may become even more relevant during episodes of financial turmoil. Many of these events have been associated with the occurrence of purely idiosyncratic shocks, like political instability, fiscal crises, or monetary indiscipline. In addition, in several of those circumstances, local authorities reacted to generalized sales of domestic assets by erecting new barriers to capital flows (e.g., pegging unilaterally their currencies and limiting the activities of foreign speculators), hence segmenting back their financial markets. Standard reduced-form ADR pricing models (e.g., Jorion and Schwartz, 1986; Foerster and Karolyi, 1999; Auguste et al., 2006) relate the dollar return for a depositary receipt to both the dollar return for the local index where the underlying security is traded and the dollar return for an index of the ADR's fully open trading venue (the U.S. equity market). The above considerations then imply that ADR dollar returns should become more sensitive to local and less sensitive to global sources of risk if normal market conditions are violated.

Testing for and explaining regime shifts in the efficiency and pricing of emerging ADR markets is therefore the subject of this work. We start by translating the arbitrage and market integration arguments made above into two linear relationships between returns for each ADR in the sample, its corresponding local stock, and domestic and global sources of risk. Our basic dataset comprises weekly U.S. dollar returns for ADRs (from their inception, whenever possible, to April 1, 2003) from issuers in 20 emerging markets: Argentina, Brazil, Chile, China, Colombia, Hong Kong, Hungary, India, Indonesia, Israel, Mexico, Peru, Philippines, Poland, Russia, South Africa, South Korea, Taiwan, Turkey, and Venezuela. Poor and time-varying liquidity in these markets (e.g., Lesmond, 2005) may induce *temporary* shifts to those linear relationships even in normal market conditions (i.e., less economically meaningful shifts, given our purposes). Further, emerging market financial crises often evolve over periods of many months or years around their official initial dates (e.g., Kaminsky et al., 2003; Kallberg et al., 2005). Thus, we search for the single, most economically and statistically significant *persistent* break over our sample period in each of the two hypothesized relationships, using the non-parametric statistical methodology of Bai et al. (1998). This approach has several desirable properties: i) it allows statistical inference on the nature of those breaks (and the estimation of confidence intervals around the estimated, endogenous break dates) with minimal restrictions on the behavior of the involved variables; ii) it exhibits adequate size and power to that end, even under the alternative hypothesis of one or more temporary structural breaks; and iii) it is unaffected by any prior bias about when a persistent break is expected to happen.

We find that the law of one price ceased to hold for many emerging depositary receipts in our sample in proximity of financial crises. Most of the estimated persistent breaks in their basic no-arbitrage relationships cluster in 1994, between 1997 and 1998, and by the end of 2001. In those circumstances, the correlation between ADR returns and the dollar returns of their perfect substitutes weakened considerably (by 54% on average), especially in Argentina, Chile, Mexico, South Africa, and Venezuela. These shifts were often accompanied by greater dependency of ADR prices on domestic equity index returns (usually by about 100%). Based on this evidence we conclude that during recent episodes of financial turmoil, the market for emerging ADRs became on average less efficient and more segmented than during more tranquil times. As interesting, however, these episodes were accompanied by either no discernible persistent deterioration or even a persistent improvement in the efficiency and/or pricing relationships of at least half of the ADRs in our database, most notably in Brazil, China, India, and Russia. This additional evidence suggests that some emerging markets may have actually benefited from the intense portfolio rebalancing activity across them reported to have taken place during those turbulent times (e.g., Disyatat and Gelos, 2001; Kallberg et al., 2005).

Our analysis goes a step further. We intend in fact to explain the occurrence, dynamics, and clustering of the estimated regime shifts. To accomplish this task, we consider a number of popular arguments in the literature (described in Section 6) to explain timing and evolution of financial crises. We devise various proxies for them, which can be grouped into six main categories: Information, trend, currency and economic environment, investors' activity, market liquidity, and expected returns. We then include the selected proxies in a Poisson model for the concentration of estimated breaks in each country in the sample. Despite our efforts, data limitations and potentially conflicting interpretations of those proxies make the resulting analysis exploratory and tentative. Given these caveats, the model's estimation indicates that emerging ADR markets are more likely to become less efficient and integrated in the presence of greater uncertainty among foreign investors and exchange rate volatility, declining expected returns and liquidity, and shrinking trade (but not prolonged stock market trends, abrupt currency devaluations, capital controls, or foreign capital flight).

This work is closely related to two recent studies. Using intraday data, Gagnon and Karolyi (2004) detect significant (albeit rarely persistent) short-term ADR price parity violations, the greater so in the presence of "excess" systematic comovement with U. S. market returns. While their analysis is static in nature, the focus of our research is on the long-term dynamics of ADR pricing and on its interaction with financial crises. Kallberg et al. (2005) find statistically significant breaks in linear relationships between the equity and currency markets of several East Asian countries during 1997 and 1998, and link those episodes to herding by foreign investors. These structural dependencies, however, do not naturally arise from normal market conditions and are not bound to hold dynamically, regardless of whether those markets were experiencing a crisis. We are also not the first to investigate the properties of emerging ADRs in proximity of episodes of financial turmoil. Among others, previous work examines the impact of selected financial crises either on the performance and valuation of selected ADRs (Mathur et al., 1998; Bin et al., 2004; Chung, 2005) or on their microstructure characteristics (Huang and Stoll, 2001) using event-study methodologies. Yet, these approaches may be affected by the analyst's prior bias for they impose an exogenous, static, and usually short event horizon — the "crisis" period — over which to evaluate ADRs' properties. The ensuing findings are time-specific by construction, hence difficult to compare to any benchmark pre- or post-crisis behavior. Our approach does not assume, but seeks to determine the nature of crisis periods by dynamically identifying endogenous periods of normal and abnormal market conditions. We do so for all emerging markets and over the longest time frame for which data is available. Lastly, we further the robustness of our analysis by relating the occurrence and clustering of normal and abnormal ADR pricing behavior to several macroeconomic and financial variables.

The paper is organized as follows. Section 2 identifies two basic efficiency and pricing relationships for ADR returns. Section 3 details the econometric approach to testing for breaks in these relationships. Section 4 describes the data employed in the analysis. We present and discuss the evidence on the estimated regime shifts in Section 5. Section 6 investigates the significance of alternative explanations for their time series and cross-country dynamics. Section 7 concludes.

2. Efficiency and pricing for ADR returns

American Depositary Receipts (ADRs) are dollar-denominated, negotiable certificates representing a pre-specified amount of a foreign company's publicly traded equity, held on deposit in the issuer's domestic market. The depositary banks pass all dividends and payments related to the underlying shares (converted into U.S. dollars) to the holders of the ADRs, who therefore bear all currency risk and pay handling fees to the custodians. ADRs allow investors to achieve international diversification while avoiding hefty local fees (up to 40 basis points, according to Velli, 1994) and trade failures, common especially in emerging markets.[2]

The holder of an ADR can redeem his certificate into the underlying shares at any point in time by simply paying a small service charge to the custodian. Conversely, new ADRs can be created at any point in time by simply depositing the appropriate number of those shares in the custodian's hands. Hence, in the absence of significant investment barriers and accounting for fluctuations in the exchange rate, ADRs and the underlying equity should be perfect substitutes (e.g., Karolyi, 1998, 2006; Gagnon and Karolyi, 2004). If r_{it}^m is the dollar return for the shares of company i in country m at time t and r_{it}^{US} is the dollar return for the corresponding depositary receipt at time t, the null hypothesis of perfect substitutability between the two securities implies that, in the relationship

$$r_{it}^{\mathrm{US}} = a_i + b_i r_{it}^m + \eta_{it}^{\mathrm{US}}, \tag{1}$$

both $a_i=0$ and $b_i=1$.[3] By construction, the validity of these return restrictions is independent of firm-specific events such as stock splits, dividend payments, capital structure decisions, or control changes. However, differential taxation rules (e.g., on the treatment of dividends), restrictions to capital mobility and foreign ownership of domestic stocks, involuntary

[2] Further details on the types of ADR programs in the U.S. and the functioning of the ADR market are in Kato et al. (1991) and Ely and Salehizadeh (2001), among others, and on the Web site of the Bank of New York (BoNY): http://www.adrbny.com/dr_edu_landing.jsp. BoNY is the world's largest depositary for ADRs.

[3] This hypothesis is equivalent to requiring that, in the absence of capital and currency controls, transaction costs, and time delays in ADR conversion, the law of one price holds for the prices of ADRs and their underlying local shares at any point in time. If P_{it}^m is the stock price of firm i in country m, P_{it}^{US} is the price of the corresponding ADR, ϕ_i^m is the conversion ratio between the two (number of local shares underlying one ADR), and S_t^m is the spot exchange rate (units of local currency per U.S. dollar), the law of one price requires that, from the perspective of a U.S. dollar-based investor, $P_{it}^{\mathrm{US}}=\phi_i^m P_{it}^m/S_t^m$ at each point in time t. Taking first differences leads to $P_{it}^{\mathrm{US}}-P_{it-1}^{US}=\phi_i^m\,(P_{it}^m/S_t^m-P_{it-1}^m/S_{t-1}^m)$. Following standard practice in the literature testing for violations of the law of one price (e.g., Auguste et al. 2006), we derive Eq. (1) by taking first differences of log prices from the no-arbitrage condition above: $\ln P_{it}^{\mathrm{US}}=\ln\phi_i^m+(\ln P_{it}^m-\ln S_t^m)$ implying that $\ln P_{it}^{\mathrm{US}}-\ln P_{it-1}^{US}=(\ln P_{it}^m-\ln P_{it-1}^m)-(\ln S_t^m-\ln S_{t-1}^m)$. This representation effectively reduces heteroskedasticity and dampens exponential growth patterns common to some stocks in our sample. Results are qualitatively similar when using arithmetic price differences and are available upon request.

dilution of ADR-holders' claims (due to SEC rules on the exercise of rights), limited liquidity, and non-synchronous trading may *temporarily* limit the efficient realignment of the dollar prices of ADRs and corresponding domestic stocks only within no-arbitrage bands such that $a_i \neq 0$ and/or $b_i \neq 1$. Most available studies (e.g., Rosenthal, 1983; Kato et al., 1991; Webb et al., 1995) conclude that the ADR market is at least weak-form efficient and that, after accounting for many of those frictions, the law of one price does indeed hold. One of the main objectives of this work is to test whether efficiency in the market for emerging ADRs is *persistently altered* during periods of financial stress. Thus, we test whether Eq. (1) experiences a regime shift in proximity of recent financial crises.[4] We analyze and discuss the importance of those frictions for the ensuing evidence in Section 6.

Established theoretical and empirical research supports the case for international diversification.[5] Hence, if the international capital markets were fully integrated and perfect capital mobility held, returns on ADRs should depend only on their covariances with global market factors relevant in their fully open trading venue (the U.S. equity market). Instead, the existence of home bias among domestic investors leads local market dynamics to affect ADR prices as well (e.g., Errunza and Losq, 1985; Alexander et al., 1987; Karolyi and Stulz, 2003). The relevance of local factors may be even greater during periods of financial stress. Many of the recent episodes of financial turmoil raging through Asia and Latin America were in fact ignited or fueled by domestic events (e.g., political uncertainty, bank failures, explosive budget deficits, or government defaults on bond payments).

In this paper we also test whether local market returns became *persistently more important* in pricing emerging ADRs around the time of financial crises. To that purpose, we employ a standard multi-factor model in which factors typically proxy for sources of domestic and U.S. market risk. The model can be written as

$$r_{it}^{\mathrm{US}} - r_{\mathrm{RF}t}^{\mathrm{US}} = b_i^m \left(r_{\mathrm{M}t}^m - r_{\mathrm{RF}t}^{\mathrm{US}} \right) + b_i^{\mathrm{US}} \left(r_{\mathrm{M}t}^{\mathrm{US}} - r_{\mathrm{RF}t}^{\mathrm{US}} \right) + e_{it}^{\mathrm{US}}, \tag{2}$$

where r_{RFt}^{US} is the U.S. risk-free rate of return at time t, $r_{\mathrm{M}t}^m$ is the dollar return on the local market portfolio of the ADR's underlying stock at time t, and r_{Mt}^{US} is the return on the U.S. market portfolio at time t. Eq. (2) is intuitively reasonable for it implies that, as a first-order approximation, ADR returns are driven by fluctuations in either the market where the ADR is traded, the market for the underlying stock, or both. Accordingly, similar specifications have been employed in several studies of cross-border listings (e.g., Jorion and Schwartz, 1986; Howe and Kelm, 1987; Karolyi, 1998; Foerster and Karolyi, 1999; Auguste et al., 2006).[6] As argued in Section 1, a significant estimate for the local (global) market coefficient b_i^m (b_i^{US}) in Eq. (2) provides evidence of relative capital market segmentation (integration). Thus, conditional on its validity as an asset pricing model, this relation allows us to test for long-lasting changes in the degree of market integration and capital mobility in the context of the emerging ADR market — i.e., to test for whether b_i^{US} and/or b_i^m experience a regime shift — in proximity of recent financial crises.

3. Testing for regime shifts

In order to test for breaks in the parameters of Eqs. (1) and (2), we adopt the statistical methodology of Bai et al. (1998). This procedure permits statistical inference about structural breaks (including the estimation of confidence interval around the break dates) with minimal distributional restrictions. Bai et al. (1998)'s non-parametric technique searches for a single, persistent break in univariate or multivariate time series models (with or without stationary regressors) and generates asymptotic confidence intervals around their estimated break dates.[7]

To that purpose, we allow for the possibility of a structural regime shift in Eqs. (1) and (2). If k is a potential break date, X_{it}^m is a $1 \times N$ vector of country- or stock-specific regressors, and β_i^m and $\Delta\beta_i^m$ are $N \times 1$ vectors of factor loadings, we specify the relation

$$y_{it}^m = X_{it}^{m\prime} \beta_i^m + d_t(k) X_{it}^{m\prime} \Delta\beta_i^m + \varepsilon_{it}^m, \tag{3}$$

where $y_{it}^m = r_{it}^{\mathrm{US}}$ or $r_{it}^{\mathrm{US}} - r_{RFt}^{\mathrm{US}}$, and $d_t(k) = 1$ if $t \geq k$ and zero otherwise. Eq. (3) is a model of full structural change since all the coefficients in β_i^m are allowed to change. Hence, the vector $\Delta\beta_i^m$ can be interpreted as the change in these coefficients after a break occurred.

We are interested in testing the null hypothesis that $\Delta\beta_i^m = 0$ at each date k. Bai et al. (1998)'s test for structural breaks to Eq. (3) considers the maximum of a Wald statistic $F_i^m(k)$ (Eq. (2.7), p. 398) constructed with OLS estimators for $B_i^m(k)$ and var (ε_i^m) (i.e., $\widehat{B}_i^m(k)$ and $\widehat{\sigma}_{im}^2$, respectively) under the alternative hypothesis of one break at date k. The estimated break date is then $\widehat{k}_i^m = \arg\max F_i^m(k)$, and is statistically significant if $F_i^m(\widehat{k}_i^m)$ is greater than its critical value at the chosen level of significance. If a statistically significant break is

[4] All our tests are run from the perspective of a dollar-based investor, since both ADR and associated local stock returns are in dollar terms. As shown in Footnote 3, fluctuations in both the local share prices and the local exchange rate versus the dollar affect the extent to which the law of one price holds. Yet, no clear predictions for their relative importance, nor any general assumption about their interaction can be formulated. Our inference is qualitatively similar when we run our tests from the perspective of local investors (i.e., using local currency returns). We explore in greater detail the role of exchange rate dynamics for the estimated regime shifts in Eq. (1) in Section 6.

[5] E.g., see Solnik (1974), Heston and Rouwenhorst (1994), and Griffin and Karolyi (1998), among others.

[6] Our inference is qualitatively similar when we replace r_{Mt}^{US} in Eq. (2) with dollar returns on a world equity index from Datastream, r_{Mt}^{W}. The correlation between r_{Mt}^{W} and r_{Mt}^{US} is 0.823 over the sample period.

[7] It is possible that those time-series models may experience more than one regime shift over the sample period. In these circumstances, Bai et al. (1998)'s methodology allows us to identify the single most (economically and statistically) significant of those breaks. Determining whether such break occurs in proximity of periods of international financial turmoil is the main objective of this research.

detected in Eq. (3), Bai et al. (1998) specify an asymptotic confidence interval for the true break date based on $\widehat{B}_i^m(k)$ and $\widehat{\sigma}_{im}^2$ (Eq. (2.21), p. 402) that does not require the restrictive assumption of normality of the residuals in Eq. (3).[8]

4. Data

Our basic dataset consists of weekly, continuously compounded, dividend-adjusted, U.S. dollar returns for all the Level II and Level III ADR stocks traded at the NYSE, NASDAQ, or AMEX (r_{it}^{US}) as of April 1, 2003, and their domestic counterparts traded in emerging financial markets (r_{it}^m). The list of all such issuers is available on the Bank of New York (BoNY)'s Web site.[9] Level II and Level III ADRs require basic to full compliance with U.S. GAAP and SEC disclosure rules. We define as emerging economies the 26 countries included in the MSCI Emerging Markets Free Index (EMF), as of April 1, 2003, plus Hong Kong.[10] We ignore stocks for which no information on their trading venue was provided by BoNY, Level 1 ADRs, and Rule 144A (also known as RADRs). Those stocks require in fact minimal or no compliance with U.S. GAAP and SEC disclosure rules, and are traded on the OTC (Over The Counter) or PORTAL (Private Offerings, Resales, and Trading through Automated Linkages) systems with much less liquidity than exchange-listed securities (e.g., Karolyi, 1998). We also ignore companies whose domestic and ADR shares have a short common trading history (100 weeks or less).

This filtering reduces our sample to joint time series of r_{it}^m and r_{it}^{US} — from the earliest available date (generally from inception) to April 1, 2003 — for 156 stocks in 20 countries. These countries are: Argentina, Brazil, Chile, China, Colombia, Hong Kong, Hungary, India, Indonesia, Israel, Mexico, Peru, Philippines, Poland, Russia, South Africa, South Korea, Taiwan, Turkey, and Venezuela. All pricing data is from Datastream. We use weekly returns to control for infrequent and non-synchronous trading, non-overlapping trading hours, and various microstructure frictions which often plague emerging ADR data at higher frequencies while still allowing for intra-month price dynamics.[11] Our dataset also includes weekly time series of three-month Treasury Bill rates (r_{RFt}^{US}) and of dollar returns of broad market indexes for each of the 20 emerging countries (r_{Mt}^m) and the U.S. (r_{Mt}^{US}), again from Datastream.

5. Estimating regime shifts

In this section we test whether the efficiency and pricing models for ADRs described in Section 2 experience a persistent regime shift in proximity of financial crises. To that purpose, we apply the methodology of Bai et al. (1998) to Eqs. (1) and (2) for each of the ADRs in our sample.[12] Tables 1 and 2 report the resulting regression coefficients from the estimation of both models at statistically significant (at the 5% level) break dates $\widehat{k}_i^m = \arg\max F_i^m(k)$, i.e., when the corresponding Wald statistic $F_i^m(\widehat{k}_i^m) \geq 11.78$, its asymptotic 95% critical value (from Table 10 in Bekaert et al., 2002). To provide a benchmark for comparison, Tables 1 and 2 also report country average estimates of Eqs. (1) and (2) for the ADRs that did not experience any such break.

The evidence in Table 1 suggests that a majority of the depositary receipts in our sample experienced a regime shift in their basic arbitrage relationships over the sample period. Specifically, we find that a statistically significant break occurs in the parameters of Eq. (1) for 81 of the 156 emerging market ADRs described in Section 4. These breaks are due on average to decreased dependence of the ADRs' dollar returns (r_{it}^{US}) on the dollar return for the underlying shares in the local market (r_{it}^m), i.e., $\widehat{\Delta b_i} < 0$ on average. Consistently, there is greater support for the null hypothesis of perfect substitutability between ADRs and local stocks when the Wald statistic $F_i^m(k)$ does not register any break: In those circumstances, estimates for b_i (a_i) in Table 1 are in fact much greater (smaller), and not statistically different from one (zero) more often than when a break occurs. Further, most of the estimated breaks took place during periods of financial turmoil. The scatter plot of estimated break dates in Fig. 1a shows that statistically significant regime shifts in Eq. (1) cluster during the events of Mexico (1994), East Asia (1997), Russia (1998), Brazil (1999), Turkey (2001), and Argentina (2001/2002), whose indicative "official" initial dates (from Table 1 in Kaminsky et al., 2003) are represented by vertical dotted lines. Accordingly, Table 1 reports that the median absolute distance between our efficiency break dates and the closest of those official crisis dates (column $|\widehat{k}_i^{m*}|$) is less than five months and as low as a few weeks for Argentina, Indonesia, Peru, and Russia.

As important, Table 3 and Fig. 1a indicate that most of the breaks in Table 1 — when implying larger violations of the law of one price (i.e., $\widehat{\Delta b_i} < 0$) — are economically significant, occurred in closest proximity of the events of Brazil, Turkey, and Argentina (but not Russia and to a lesser extent the East Asian crisis, in columns $|\widehat{k}_{iE}^{m*}|$ and $\widehat{\Delta b_i} < 0$ of Table 3), and especially concern ADRs from Argentina, Chile, Mexico, South Africa, and Venezuela.[13] In particular, the average relative deterioration of the law of one price (i.e., with respect to pre-break b_i, whose mean is 0.95) i) amounts to a considerable 54% ($\overline{\widehat{\Delta b_i}} = -0.49$ in Table 3), ii) occurs within (a median of) only 15 weeks from the closest official initial crisis date (i.e., economically close to those crises' observed chronological

[8] Albeit constructed using limit theorems, these tests display satisfactory finite-sample properties: According to Bai et al. (1998) and Bekaert et al. (2002), they perform adequately, in terms of both size and power, under the null hypothesis of no break and the alternative hypothesis either of a single persistent break in the mean of the dependent variable or of persistent structural breaks in the coefficients of exogenous regressors, as in Eq. (3). Further Monte Carlo analysis specific to our empirical setting (available on request) confirms these findings, even under the null hypothesis of either a single or multiple temporary breaks of various duration and intensity to the parameters of Eq. (3).

[9] Http://www.adrbny.com/dr_directory.jsp.

[10] Http://www.msci.com/licensing/derive.html. We add Hong Kong to the MSCI list because of the presence of many Chinese companies in that stock exchange.

[11] E.g., see Harvey (1995), Aggarwal et al. (1999), and Bacidore and Sofianos (2002). Auguste et al. (2006) take weekly averages of daily ADR premiums or discounts for Argentine stocks. Unreported results based on lower frequencies (e.g., monthly) are qualitatively similar.

[12] The correlations between r_{Mt}^{US} and each of the local aggregate return series r_{Mt}^m are relatively low, ranging from a minimum of 0.07 versus India to a maximum of 0.37 versus Mexico. Therefore, when we estimate Eq. (2), we do not orthogonalize the regressors r_{Mt}^m, as is instead done in Auguste et al. (2006).

[13] Consistently, Hunter (2005) finds evidence of deteriorating efficiency in the ADR markets of Argentina, Chile, and Mexico during crisis periods.

Table 1
Efficiency tests

		One break								No break			
	N	N^*	$\overline{\widehat{k}_i^{m\pm}}$	$\lvert\widehat{k}_i^{m}{}^*\rvert$	$\overline{R_a^2}$	$\overline{\widehat{a}_i}$	$\overline{\widehat{b}_i}$	$\overline{\widehat{\Delta a_i}}$	$\overline{\widehat{\Delta b_i}}$	N^o	$\overline{R_a^2}$	$\overline{\widehat{a}_i}$	$\overline{\widehat{b}_i}$
Total	156	81	63 (58)	18 (115)	63.35%	−0.0017 (4)	0.7124 (73,21)	0.0006 (5)	−0.0236 (81)	75	72.14%	−0.0009 (1)	0.8734 (73,28)
Argentina	13	9	33 (19)	7 (12)	54.49%	−0.0019 (0)	0.7754 (8,5)	0.0017 (0)	−0.0205 (9)	4	38.06%	−0.0039 (0)	0.6175 (3,1)
Brazil	30	12	47 (26)	16 (15)	75.40%	0.0003 (0)	0.7726 (11,3)	−0.0016 (0)	0.1245 (12)	18	84.32%	−0.0002 (0)	0.9235 (18,6)
Chile	19	10	105 (84)	18 (60)	50.34%	−0.0008 (0)	0.6673 (7,4)	−0.0002 (1)	−0.0899 (10)	9	77.50%	−0.0004 (0)	0.9205 (9,4)
China	14	8	54 (36)	25 (7)	81.97%	−0.0036 (0)	0.5884 (7,0)	0.0035 (0)	0.2853 (8)	6	85.29%	0.0002 (0)	0.9112 (6,2)
Colombia	1	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	1	4.69%	−0.0029 (0)	0.2753 (1,0)
Hong Kong	6	1	42 (n.a.)	12 (n.a.)	59.89%	−0.0113 (0)	0.5876 (1,0)	0.0115 (0)	0.4134 (1)	5	65.33%	−0.0017 (0)	0.8516 (5,1)
Hungary	1	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	1	82.65%	−0.00003 (0)	0.9426 (1,0)
India	7	2	20 (8)	14 (1)	68.26%	0.0143 (1)	0.6578 (1,1)	−0.0161 (1)	0.1764 (2)	5	56.86%	−0.0017 (0)	0.8151 (5,2)
Indonesia	2	2	61 (27)	5 (6)	81.29%	−0.0034 (0)	0.6358 (2,0)	0.0038 (0)	0.2296 (2)	0	n.a.	n.a.	n.a.
Israel	7	3	65 (39)	24 (38)	55.33%	0.0021 (0)	0.4534 (3,0)	−0.0023 (0)	0.4275 (3)	4	80.36%	−0.0002 (0)	0.9698 (4,3)
Mexico	27	16	76 (61)	29 (77)	68.12%	−0.0018 (2)	0.7967 (16,4)	−0.0001 (2)	−0.1991 (16)	11	74.65%	−0.0001 (0)	0.9218 (11,5)
Peru	2	1	32 (n.a.)	4 (n.a.)	43.59%	−0.0014 (0)	0.8168 (1,0)	0.0044 (0)	−0.5168 (1)	1	−0.61%	−0.0147 (1)	−0.0023 (0,0)
Philippines	1	1	298 (n.a.)	27 (n.a.)	55.32%	0.0012 (0)	0.7092 (1,0)	−0.0018 (0)	0.1880 (1)	0	n.a.	n.a.	n.a.
Poland	1	1	12 (n.a.)	8 (n.a.)	51.65%	−0.0053 (0)	0.9165 (1,1)	−0.0134 (0)	−1.1045 (1)	0	n.a.	n.a.	n.a.
Russia	3	3	17 (6)	8 (0)	45.29%	−0.0206 (1)	0.5409 (3,1)	0.0260 (1)	0.0487 (3)	0	n.a.	n.a.	n.a.
S. Africa	8	6	57(44)	20 (361)	64.75%	−0.0009 (0)	0.8259 (5,2)	0.0022 (0)	−0.1890 (6)	2	63.33%	−0.0005 (0)	0.8087 (2,0)
S. Korea	6	4	81 (46)	25 (12)	50.37%	−0.0027 (0)	0.5695 (4,0)	−0.0009 (0)	0.0416 (4)	2	71.73%	−0.0008 (0)	0.8780 (2,0)
Taiwan	5	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	5	70.96%	0.0001 (0)	0.9718 (5,4)
Turkey	1	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	1	66.39%	−0.0060 (0)	0.8639 (1,0)
Venezuela	2	2	44 (57)	23 (29)	42.82%	0.0025 (0)	0.7479 (2,0)	−0.0132 (0)	−0.8359 (2)	0	n.a.	n.a.	n.a.

This table reports mean OLS estimates for Eq. (1) when it either does or does not experience a statistically significant break (at the 5% level) for each country in the sample and, in brackets, the number of ADRs for which they are (not) different from zero (one) at the 5% level. R_a^2 is the adjusted R^2; N is the number of ADRs; N^* (N^o) is the subset where date $\widehat{k}_i^m$=arg max F_i^m (k) and F_i^m $(\widehat{k}_i^m)\geq 11.78$ (F_i^m $(\widehat{k}_i^m)< 11.78$); $\widehat{k}_i^{m\pm}$ is the mean width of its confidence interval; $\lvert\widehat{k}_i^{m*}\rvert$ is its median absolute distance from the closest initial financial crisis date (from Table 1 in Kaminsky et al., 2003), both in weeks with subscript standard deviation.

evolution (e.g., Kaminsky et al., 2003; Kallberg et al., 2005)), and iii) entails more than 70% of the ADRs from the 7 countries whose average shift $\overline{\widehat{\Delta b_i}}$ is negative in Table 1. Those countries' average confidence intervals around their estimated break dates (column $\overline{\widehat{k}_i^{m\pm}}$ in Table 1) are generally tighter as well (except in Chile), ranging between 3 and 18 months. Lastly, efficiency breaks for Argentine, South African, and Venezuelan ADRs are concentrated in the latter part of the sample, while substitutability between depositary receipts from Mexican and Chilean issuers and their underlying local stocks declined throughout the 1990s (see Fig. 1a). We conclude from this evidence that most of the recent financial crises involving emerging economies were accompanied by lower efficiency in their ADR markets.

Interestingly, the country- and crisis-level evidence in Tables 1 and 3 also reveal that many estimated realignments are in the direction of the law of one price: The average shift $\overline{\widehat{\Delta b_i}}$ is positive (0.41 in Table 3, pushing b_i toward one from a pre-break mean of 0.49) for 9 of the 16 countries (and for 42 of the 81 companies) where we find statistically significant breaks in Eq. (1), most notably in Brazil, China, India, and Russia. In those 9 countries, about 81% of the ADRs whose estimated break is statistically significant experience an improvement in efficiency. However, these improvements are on average of smaller magnitude and occur within (a median of) 25 weeks from official initial crisis dates, i.e., 10 weeks later than deteriorations. The trend for greater economic and financial integration, better liquidity, and increasing interest of foreign investors exhibited by most emerging markets in the last two decades (e.g., see Bekaert et al., 2002; Lesmond, 2005) may have contributed to the greater efficiency of their ADRs over our sample period. In addition, extant studies (e.g., Disyatat and Gelos, 2001; Kallberg et al., 2005) find that the portfolio rebalancing activity of foreign investors across emerging markets greatly intensified during those crises and that some of these markets were net recipients of funds.

In conjunction with the above results, this evidence suggests that portfolio rebalancing across emerging markets may have improved the efficiency of some of them and lessened the efficiency of others. We investigate more directly the role of the trading activity of foreign investors for the occurrence and clustering of efficiency breaks in proximity of periods of financial turmoil in Section 6.

According to Tables 2 and 3 and Fig. 1b, recent periods of financial turmoil were also associated with regime shifts in the ADR pricing model of Eq. (2). Table 2 reports that 79 ADRs in our sample experience a statistically significant break in the relationship between their excess dollar returns and domestic and global sources of risk. In correspondence with these breaks, the estimated $\Delta\widehat{b}_i^m$ is positive on average — indicating greater covariance of ADR returns with local risk factors — in 9 of the 15 countries, and 45 of the ADRs in our sample for which at least one regime shift in Eq. (2) is found (for a median relative increase of 109% with respect to the pre-break $\widehat{b}_i^m$). Similarly, $\Delta\widehat{b}_i^{\mathrm{US}}$ is on average negative and large — indicating lower covariance of ADR returns with world risk factors — for 8 countries and 44 ADRs in Table 2 (for a median relative decline of 99% with respect to $\widehat{b}_i^{\mathrm{US}}$).[14] According to Table 3 (columns $\Delta\widehat{b}_i^m>0$ and $\Delta\widehat{b}_i^{\mathrm{US}}>0$), the persistent shifts in the sensitivity of those ADRs' returns to domestic and global sources of risk are economically significant as well, averaging 0.74 and −0.66 versus pre-break means of 0.58 and 0.60, respectively. In contrast to

[14] Table 2 also indicates that both the average confidence intervals around the estimated pricing break dates $\widehat{k}_i^m$ plotted in Fig. 1b, $\overline{\widehat{k}_i^{m\pm}}$, and their median absolute distances from the closest initial crisis date in Kaminsky et al. (2003), $\lvert\widehat{k}_i^{m*}\rvert$, are generally larger (albeit more homogeneously so) than those for the efficiency break dates in Fig. 1a (reported in Table 1).

Table 2
Pricing tests

		One break								No break			
	N	N^*	$\overline{\widehat{k}_i^{m\pm}}$	$\lvert\widehat{k}_{i^m}^*\rvert$	$\overline{R_a^2}$	$\overline{\widehat{b}_i^m}$	$\overline{\widehat{b}_i^{US}}$	$\overline{\Delta\widehat{b}_i^m}$	$\overline{\Delta\widehat{b}_i^{US}}$	N^o	$\overline{R_a^2}$	$\overline{\widehat{b}_i^m}$	$\overline{\widehat{b}_i^{US}}$
Total	156	79	71 (53)	25 (77)	35.69%	0.7538 (71)	0.2929 (28)	0.1274 (67)	−0.0975 (13)	77	30.84%	0.9498 (74)	0.2654 (34)
Argentina	13	9	55 (50)	12 (16)	37.15%	0.8061 (9)	0.1705 (2)	0.0389 (8)	0.5965 (5)	4	8.91%	0.5388 (4)	0.1182 (1)
Brazil	30	13	73 (52)	22 (18)	48.90%	0.8737 (11)	0.3546 (4)	0.2853 (12)	−0.2850 (4)	17	39.99%	0.9734 (17)	0.2112 (7)
Chile	19	11	69 (46)	28 (41)	25.94%	0.6647 (8)	0.4580 (5)	0.1943 (9)	−0.2861 (5)	8	35.29%	1.1769 (8)	0.0890 (0)
China	14	8	54 (36)	24 (9)	24.63%	0.5838 (7)	−0.0983 (2)	0.1151 (7)	0.1865 (3)	6	24.77%	0.6409 (4)	0.2487(5)
Colombia	1	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	1	5.93%	0.5264 (1)	0.3422 (1)
Hong Kong	6	2	85 (27)	16 (17)	18.45%	0.1857 (1)	−0.0477 (0)	0.6063 (2)	0.4988 (1)	4	15.97%	1.0619 (4)	0.3049 (1)
Hungary	1	1	34 (n.a.)	26 (n.a.)	70.87%	0.7607 (1)	0.1067 (0)	0.5046 (1)	0.1558 (0)	0	n.a.	n.a.	n.a.
India	7	2	51 (24)	31 (26)	31.25%	1.2523 (2)	−0.8646 (1)	−0.2916 (1)	1.1265 (1)	5	30.39%	1.2401 (5)	0.5629 (2)
Indonesia	2	2	59 (35)	8 (1)	59.17%	0.4949 (2)	0.2574 (0)	0.3809 (2)	0.1427 (0)	0	n.a.	n.a.	n.a.
Israel	7	3	79 (47)	29 (25)	28.59%	0.5593 (3)	0.6085 (1)	0.5356 (2)	−0.4979 (1)	4	29.23%	1.0782 (4)	0.3524 (3)
Mexico	27	17	69 (38)	19 (21)	38.43%	0.8675 (16)	0.4373 (7)	−0.0065 (15)	−0.3910 (6)	10	30.79%	0.8522 (9)	0.2000 (2)
Peru	2	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	2	18.55%	1.2486 (2)	0.2589 (1)
Philippines	1	1	62 (n.a.)	49 (n.a.)	43.24%	0.8409 (1)	1.0825 (1)	−0.1150 (0)	−1.0442 (1)	0	n.a.	n.a.	n.a.
Poland	1	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	1	4.67%	1.2698 (1)	0.2328 (0)
Russia	3	2	13 (1)	8 (1)	47.25%	0.2370 (2)	1.6028 (1)	0.9030 (2)	−1.1883 (1)	1	38.52%	0.8666 (1)	0.6097 (1)
S. Africa	8	5	178 (75)	61 (258)	17.84%	0.7814 (5)	−0.3712 (2)	−0.1064 (3)	0.2516 (2)	3	14.64%	0.7245 (3)	−0.3101 (3)
S. Korea	6	2	64 (71)	31 (5)	28.78%	0.7858 (2)	0.9817 (1)	−0.2345 (2)	−0.7148 (0)	4	32.52%	0.8847 (4)	0.4061 (2)
Taiwan	5	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	5	55.16%	1.1398 (5)	0.7219 (5)
Turkey	1	0	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	1	48.74%	0.8499 (1)	0.2670 (0)
Venezuela	2	1	16 (n.a.)	8 (n.a.)	55.29%	1.1118 (1)	0.2958 (1)	−1.1376 (1)	0.3283 (0)	1	2.53%	0.4710 (1)	0.4596 (0)

This table reports mean OLS estimates for Eq. (2) when it either does or does not experience a statistically significant break (at the 5% level) for each country in the sample and, in brackets, the number of ADRs for which they are (not) different from zero (one) at the 5% level. R_a^2 is the adjusted R^2; N is the number of ADRs; N^* (N^o) is the subset where date $\widehat{k}_i^m = \arg\max F_i^m(k)$ and $F_i^m(\widehat{k}_i^m) \geq 11.78$ ($F_i^m(\widehat{k}_i^m) < 11.78$); $\overline{\widehat{k}_i^{m\pm}}$ is the mean width of its confidence interval; $\lvert\widehat{k}_i^{m}{}^*\rvert$ is its median absolute distance from the closest initial financial crisis date (from Table 1 in Kaminsky et al., 2003), both in weeks with subscript standard deviation.

Fig. 1a, Fig. 1b further reveals that pricing breaks tend to cluster in geographical patterns, especially in Latin America. For example, most shifts to Eq. (2) took place in late 1994 for Mexican issuers, in 1999 for Brazilian issuers, and in 2001 for Argentine issuers. Nonetheless, consistent with visual inspection of Fig. 1b, Table 3 also shows that the estimated pricing break dates of Table 2 are close to the official initial crisis dates in Kaminsky et al. (2003) — between 7 and 42 weeks in column $\lvert\widehat{k}_{iP}^{m}{}^*\rvert$ — although generally less so than the estimated efficiency breaks of Table 1 (column $\lvert\widehat{k}_{iE}^{m}{}^*\rvert$).

When compared to the benchmark estimates of Eq. (2) in Table 2 (i.e., in the absence of a break point), this evidence suggests that with few exceptions (e.g., Argentina and Venezuela), most recent episodes of financial turmoil entailed more intense generation of information in the less liquid local markets of emerging economies than in the U.S. stock market. As important, Table 3 further shows that, in proximity of global financial turmoil, half of the emerging ADRs whose returns became significantly more sensitive to domestic factors and less sensitive to global factors ($\Delta\widehat{b}_i^m > 0$ & $\Delta\widehat{b}_i^{US} < 0$) — and as many as 75% to 100% of them during the Mexican and East Asian crises — were also experiencing more severe arbitrage violations ($\widehat{\Delta b_i} < 0$). We conclude that during recent financial crises not only was the market for emerging ADRs generally less efficient but also less integrated, as argued in Sections 1 and 2.

6. Explaining regime shifts

The evidence presented in the previous section can be interpreted as supportive of the notion (postulated in Sections 1 and 2) that protracted periods of economic and financial instability may translate into important, persistent shifts in both the efficiency of emerging ADR markets and the relevance of the pricing factors driving their dynamics. In the remainder of the paper, we attempt to explain why these shifts took place in proximity of the many crises that occurred in the past decade or so. To that purpose, we introduce two new variables. The first is E_t^m, the number of companies in country m whose ADRs are experiencing a statistically significant permanent break in the efficiency relationship of Eq. (1) at time t. Hence, E_t^m measures the extent of concentration and persistence of efficiency break events across all ADRs in country m at week t. Similarly, we define P_t^m as the number of companies in country m whose ADRs are experiencing a statistically significant break in the pricing relationship of Eq. (2) at week t.[15] Our objective is to explain the behavior of E_t^m and P_t^m for each of the countries in our sample.

As mentioned in Section 1, there is an increasing body of research studying why financial crises occur. As such, this literature can help shed light on the clustering and dynamics of efficiency and pricing shifts for emerging ADRs in proximity of episodes of financial turmoil. We consider a number of proxies for those arguments, grouped into six categories. The first group contains financial variables related to *information*. It is often argued that information asymmetry and heterogeneity among investors with respect to asset valuations contribute to the occurrence, severity, and propagation of financial crises.[16]

[15] Specifically, E_t^m (P_t^m) is computed as the number of ADRs for country m for which t is greater than or equal to $\widehat{k}_i^{m-}$, the lower band of their confidence interval around their estimated efficiency (pricing) break date $\widehat{k}_i^m$.

[16] Pasquariello (2007) reviews this literature.

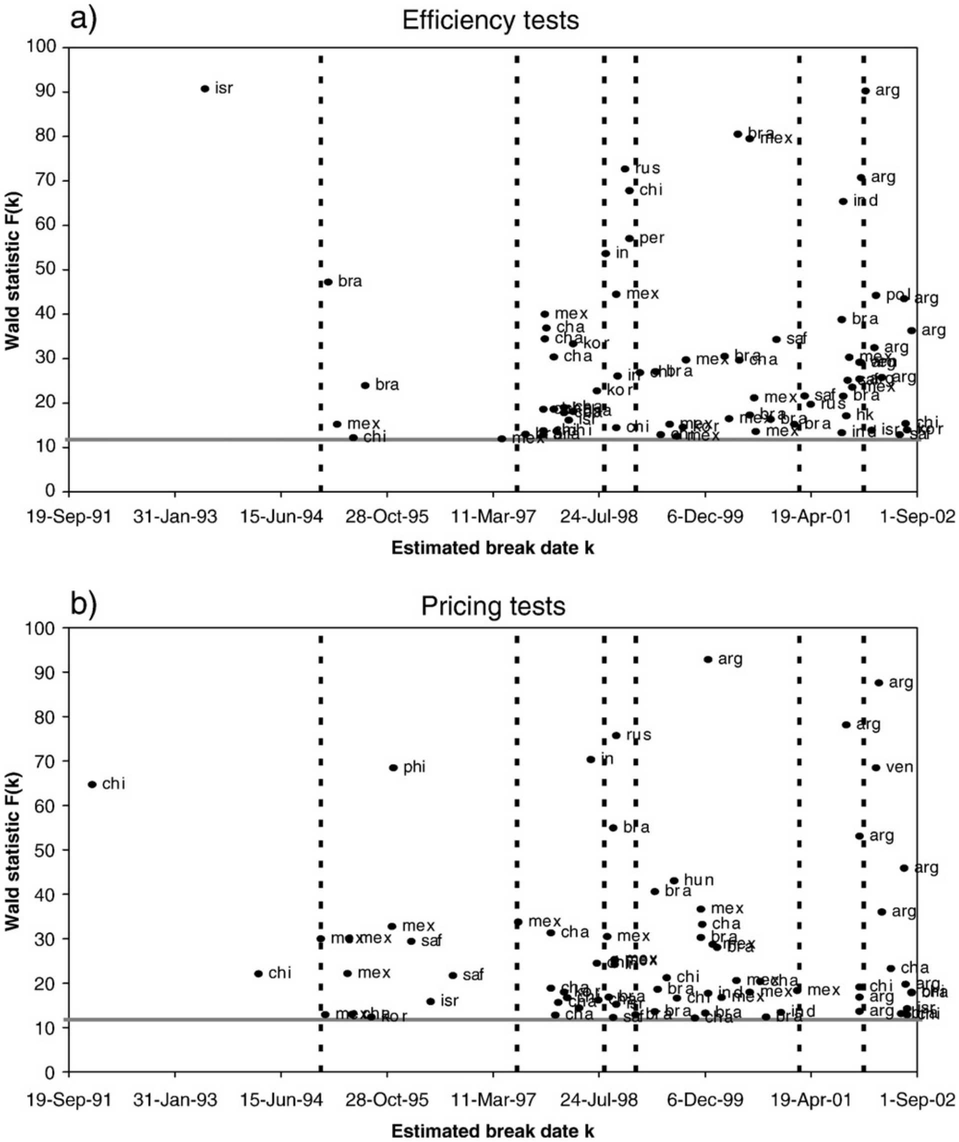

Fig. 1. Scatter plots of statistically significant regime shifts These figures are scatter plots of the maximum of the Wald statistic $F_i^m(k)$ versus the corresponding estimated break date, $\widehat{k}_i^m$, for the efficiency model of Eq. (1), in (a), and the pricing model of Eq. (2), in (b), for the subset of the 156 ADRs in our sample whose $\max_k F_i^m(k)$ is greater than its 5% significance level, 11.78 (the straight line). The figures also display the corresponding country label for those ADRs experiencing statistically significant breaks, as well as the initial dates for recent episodes of international financial turmoil (from Table 1 in Kaminsky et al., 2003) as vertical dotted lines: Mexico (December 20, 1994); Thailand (July 2, 1997), Russia (August 18, 1998), Brazil (January 13, 1999), Turkey (February 22, 2001), and Argentina (December 23, 2001).

Time-varying uncertainty and disagreement may also explain the adjustment in the covariance of ADR returns with local and global risk factors. We use domestic and U.S. stock return volatility — σ_t^m and σ_t^{US}, computed as the standard deviation of weekly equity index returns in local currency (r_{Mt}^m and r_{Mt}^{US} from Datastream) over the past 100 weeks — as (albeit imperfect) ex-post proxies for the corresponding market's information environment at time t (e.g., Lang and Lundholm, 1993; Leuz and Verrecchia, 2000; Lim, 2001; Zhang, 2006).

Generalized purchases or sales of assets within a market (e.g., motivated by imitation, injections and redemptions of cash in mutual funds, the activity of momentum traders, or bubbles) may also alter the link between ADR prices and their underlying assets. For example, Kyle and Xiong (2001) and Yuan (2005) argue that financial crises are more likely to occur when financial constraints induce rational speculators to trade regardless of fundamentals. This argument implies that violations of the arbitrage relation of Eq. (1) or shifts in the relative importance of local and global sources of risk for ADR returns in Eq. (2) should be more likely during a bear market or a correction, i.e., when those financial constraints are more binding for investors. The second group of variables therefore proxies for prolonged, asymmetric *trends* in the aggregate

Table 3
Efficiency and pricing breaks in proximity of crises

All events				% of N_E^* where				% of N_P^* where			% of N_E^* where
	Crisis date	N_E^*	$\lvert\hat{k}_{iE}^{m*}\rvert$	$\widehat{\Delta b_i}<0$	$\widehat{\Delta b_i}>0$	N_P^*	$\lvert\hat{k}_{iP}^{m*}\rvert$	$\Delta\hat{b}_i^m>0$	$\Delta\hat{b}_i^{US}<0$	$\Delta\hat{b}_i^m>0$ & $\Delta\hat{b}_i^{US}<0$	$\widehat{\Delta b_i}<0$ & $\Delta\hat{b}_i^m>0$ or $\Delta\hat{b}_i^{US}<0$
	n.a.	81	18	48% (−0.49)	52% (0.41)	79	25	57% (0.74)	56% (−0.66)	29% (0.83, −0.83)	49% (−0.45, −0.12, −0.83)
Mexico	20-Dec-94	9	78	44% (−0.35)	56% (0.65)	13	42	62% (0.46)	92% (−0.73)	54% (0.46, −0.70)	75% (−0.42, −0.29, −0.45)
Thailand	2-Jul-97	11	19	27% (−0.34)	73 (0.24)	7	26	71% (0.47)	43% (−0.46)	14% (0.06, −1.13)	100% (−0.34, −0.56, −0.42)
Russia	18-Aug-98	11	9	0%	100 (0.30)	16	7	63% (0.55)	63% (−0.59)	31% (0.55, −0.81)	0%
Brazil	13-Jan-99	10	15	50% (−0.42)	50% (0.33)	16	42	31% (0.60)	44% (−0.58)	6% (0.03, −1.27)	40% (−0.28, −0.63, −0.40)
Turkey	22-Feb-01	15	33	73% (−0.45)	27% (0.54)	7	26	57% (0.87)	86% (−0.55)	43% (1.10, −0.74)	55% (−0.53,0.19, −0.64)
Argentina	23-Dec-01	25	12	64% (−0.61)	36% (0.55)	20	26	65% (1.17)	30% (−0.92)	30% (1.60, −0.92)	31% (−0.51,0.06, −0.67)

This table reports summary statistics on the interaction between statistically significant breaks in the efficiency relationship of Eq. (1) and those in the pricing relationship of Eq. (2) in proximity of financial crises and, in brackets, the average estimated shifts in the corresponding coefficients. In particular, N_E^* and N_P^* are the number of ADRs for which a statistically significant efficiency and pricing break was observed over our sample period, respectively — i.e., where date $\hat{k}_i^m$ = arg max $F_i^m(k)$ and $F_i^m(\hat{k}_i^m) \geq 11.78$ for Eqs. (1) and (2). Each statistically significant break is then assigned to the closest financial crisis occurring over our sample period (Mexico, Thailand, Russia, Brazil, Turkey, or Argentina), i.e., to the crisis whose initial date (from Table 1 in Kaminsky et al., 2003) is the closest (in absolute terms) to the corresponding break date $\hat{k}_i^m$. Hence, $|\hat{k}_{iE}^{m*}|$ and $|\hat{k}_{iP}^{m*}|$ are the resulting median absolute distances (in weeks) of those efficiency and pricing break dates, respectively, from the corresponding crises' initial dates.

domestic and U.S. stock return series with two sets of dummy measures d_t^{m+} and d_t^{m-} (d_t^{US+} and d_t^{US-}) equal to one if the sign of r_{Mt}^m (r_{Mt}^{US}) over the current and each of the previous two weeks is positive or negative, respectively, and zero otherwise.

Our third group of variables is meant to capture the dynamics of the local and U.S. *currency and economic environment*. Many recent financial crises were ignited by pressures on the domestic currency before propagating to the local equity markets (e.g., Bailey et al., 2002; Kallberg et al., 2005; Pavlova and Rigobon, 2007). Instability in the exchange rate market and asymmetric currency dynamics may affect both Eqs. (1) and (2) either directly (when altering the terms of trade between ADRs and underlying assets) or indirectly (by virtue of their impact on the domestic and U.S. stock markets). We account for this possibility by computing two types of variables: i) the standard deviation of weekly returns for the exchange rate between the domestic currency for country m and the dollar (r_{mt}^{FX} from Datastream) over the past 100 weeks, σ_{mt}^{FX}, to measure the stability of that country's foreign exchange market (as in Bekaert and Harvey, 2000; Bekaert et al., 2002), and ii) an additional set of dummies equal to one if the local currency has been depreciating (d_{mt}^{FX+}) or appreciating (d_{mt}^{FX-}) versus the dollar over the current and each of the previous two weeks, and zero otherwise. We also consider the potential impact of both the size of the trade sector of a country and its macroeconomic environment on the efficiency and pricing of emerging ADRs. Either of these conditions may affect that country's sensitivity to fluctuations of the exchange rate and ability to fulfill its obligations toward creditors, thus ultimately investors' interest in its capital markets and currency. Trade linkages may also provide a channel for the occurrence and propagation of financial crises (e.g., Glick and Rose, 1999; Dasgupta, 2000). As in Bekaert et al. (2002), we proxy for the extent of those linkages with the monthly ratio of nominal exports plus imports divided by nominal GDP (in local currency, from IFS), T_t^m. In the spirit of Erb et al. (1996) and Bekaert and Harvey (2000), we use the foreign currency sovereign credit rating history from Standard & Poors to construct monthly series C_t^m of numeric rating shifts (from a high of 26 for AAA to a low of 1 for Default) as an (albeit imprecise) measure of the ex-ante credibility of country m's economic policies.

Over the past two decades, most of the emerging economies in our sample experienced significant regulatory changes facilitating the access of foreign investors to domestic assets (e.g., Bekaert and Harvey, 1995, 1997). Speculative trading, especially by foreign investors in emerging economies, may make their capital markets more efficient by augmenting the information content of prices (Grossman and Stiglitz, 1980; Bekaert and Harvey, 2000). Thus, foreign speculators may improve the efficiency and pricing of the market for emerging ADRs by enhancing its informational and allocational properties. Foreign investors have also been accused of destabilizing emerging capital markets (e.g., Kaminsky et al., 2001, 2004). In those circumstances, short-term trading by foreign speculators and the portfolio rebalancing activity of institutional investors would negatively affect some of these markets' ability to process information (hence worsening their efficiency and pricing) while instead potentially benefiting others (as mentioned in Section 5). Following Tesar and Werner (1995b), we proxy for *signed net trading activity by foreign investors* in each emerging equity market m in our sample by computing a fourth set of variables, F_t^m, as the net monthly equity capital flows from the U.S. for that country in week t (in millions of dollars, from the U.S. Treasury Bulletin).

Our fifth group of variables tries to measure the *liquidity* in each of the markets where ADRs and underlying stocks are traded. There is much anecdotal and empirical evidence that local market liquidity deteriorates during financial crises, stirring speculators toward more liquid trading venues, limiting arbitrage activities, and moving prices away from fundamentals (e.g., Chowdhry and Nanda, 1991; Summers, 2000; Lesmond, 2005). These phenomena may then affect the extent to which ADR prices departed from basic no-arbitrage relations or became more (less) sensitive to global (local) risk factors in proximity of periods of financial turmoil (e.g., Rabinovitch et al., 2003). As in Bekaert et al. (2002), we proxy for the liquidity of the U.S. (L_t^{US}) and each emerging (L_t^m) equity market in our database by computing the ratio of weekly volume of trading to weekly market capitalization (both in local currency, from Datastream).

Lastly, we consider the possibility that relative changes in *expected returns* influence the efficiency and pricing of ADRs, e.g., by swaying investors' expectations and preferences for local versus U.S. equities. According to Bekaert and Harvey (2000), changes in dividend yields are effective measures for changes in expected returns of emerging capital markets. Hence, we use weekly time series of annualized dividend yield levels from Datastream to define two new sets of variables, D_t^m and D_t^{US}, in each of the emerging markets in our sample and in the U.S. stock market.[17]

6.1. The analysis

We now test whether the various proxies devised above can explain the occurrence and clustering of the persistent efficiency and pricing breaks reported in Tables 1 and 2 and displayed in Fig. 1. To accomplish this task, the simplest approach would be to use multivariate linear regressions. However, because the breaks are (by definition) unique events and concentrate over relatively short periods of time, the discrete dependent variables E_t^m and P_t^m show a preponderance of zeros and small values. According to the econometric literature (e.g., Greene, 1997), the Poisson regression model is a better specification to account for these properties of the data. This model assumes that each discrete dependent variable Y_t^m (either E_t^m or P_t^m) is drawn from a Poisson distribution with parameter λ_t^m, such that

$$\Pr(Y_t^m = i) = \frac{e^{-\lambda_t^m}(\lambda_t^m)^i}{i!}, \tag{4}$$

where i is the number of ADRs in country m experiencing a break in Eq. (1), when $Y_t^m = E_t^m$, or in Eq. (2), when $Y_t^m = P_t^m$.[18] We impose that the parameter λ_t^m is related to our six groups of regressors in terms of the following log-linear model:

$$\ln\lambda_t^m = \gamma_1\sigma_t^m + \gamma_2\sigma_t^{US} + \gamma_3 d_t^{m+} + \gamma_4 d_t^{m-} + \gamma_5 d_t^{US+} + \gamma_6 d_t^{US-} + \gamma_7\sigma_{mt}^{FX} + \gamma_8 d_{mt}^{FX+} + \gamma_9 d_{mt}^{FX-} + \gamma_{10}T_t^m + \gamma_{11}C_t^m + \gamma_{12}F_t^m + \gamma_{13}L_t^m + \gamma_{14}L_t^{US} + \gamma_{15}D_t^m + \gamma_{16}D_t^{US}. \tag{5}$$

We estimate the parameters of Eq. (5) by maximum likelihood, report the results in Table 4 for $Y_t^m = E_t^m$ and in Table 5 for $Y_t^m = P_t^m$ for each emerging country m in our sample, and then discuss our evidence separately for countries experiencing either a deterioration or an improvement in their efficiency and/or pricing relationships.[19]

Before proceeding, an important premise is necessary. Despite their popularity, there is little work exploring the properties of most of the regressors in Eq. (5) or discriminating among their numerous, often potentially conflicting interpretations. As such, the analysis that follows is of an exploratory nature and any inference based upon it should be interpreted with caution. With these caveats in mind, our model appears to perform well for both sets of dependent variables E_t^m and P_t^m. The Poisson model produces no obvious counterpart to the R^2 of linear regressions. Yet, two alternative measures of goodness of fit based on standardized residuals (R_P^2) and log-deviances (R_D^2) are mostly greater than 90%, while the chi-squared statistic (χ^2) rejects the null hypothesis that the slopes in Eq. (5) are all zero for all countries but Hong Kong and Poland in Table 4 and South Korea in Table 5. Further, our evidence suggests that information asymmetry, domestic and U.S. economic conditions, and changes in relative liquidity and expected returns are the most relevant factors in explaining the clustering of efficiency and pricing breaks over the last decade.[20]

To begin with, we find that the likelihood of an emerging ADR market to become less efficient is negatively related to domestic equity volatility and positively related to U.S. stock market volatility. Specifically, we find statistically significant (often at the 1% level) $\widehat{\gamma}_1<0$ and/or $\widehat{\gamma}_2>0$ in Eq. (5) for most of those countries (e.g., Argentina, Chile, Mexico, Peru, South Africa, and Venezuela, but not Poland) whose ADR markets experienced a deterioration in their efficiency relationships over the sample period — i.e., where on average $\widehat{\Delta b_i}<0$ in Eq. (1) according to Table 1. Not surprisingly, in those same countries the greater the fluctuations in the domestic currency the more likely are ADR returns to violate the law of one price: $\widehat{\gamma}_7>0$ in Eq. (5) for E_t^m. Along those lines, the coefficient γ_7 is instead negative and significant for China, Hong Kong, India, and Israel (but not Brazil and Russia), whose ADR markets were found to become more (rather than less) efficient during the 1990s (again in Table 1).[21] However, the impact of the

[17] As in Bekaert and Harvey (2000), the weekly emerging market dividend yield series (D_t^m) are highly persistent, with an average first-order autocorrelation coefficient of 0.974. Dividend yields in the U.S. (D_t^{US}) are characterized by a pronounced downward trend over the sample period, which can be in part ascribed to the increasing popularity of alternative payout policies (e.g., stock repurchases or stock rights) by U.S. corporations. No such trends can be identified for the dividend yield series D_t^m.

[18] By construction, both time series E_t^m and P_t^m (as well as their first differences) display high positive serial correlation, thus violating the assumption that event occurrences generated by a Poisson point process in successive time intervals are independent. Unfortunately, alternative time series regression models for counts with those properties are neither available nor well developed yet (e.g., see the surveys by Cameron and Trivedi, 1996; Greene, 1997, Chapter 19). The results from multiple linear regressions (available on request) are nonetheless qualitatively similar.

[19] For each country m, the sample over which Eqs. (4) and (5) are estimated starts on the first potential break date for the ADR with the earliest return data available (t_0 in Tables 4 and 5), hence is made of no less than $N=247$ weeks (India) and up to 1, 210 weeks (South Africa). The specification of Eq. (5) has no constant term to avoid the possibility of collinear regressors in the estimation.

[20] The model of Eq. (5) does not explicitly account for the possibility that the total number of ADRs in a country at each point in time may itself be affected by that country's degree of financial development and vulnerability to crisis events. We discuss in this section the impact of time-varying local market openness on the country-level evidence presented in Section 5. At the firm level, issuance of ADRs has been motivated by liquidity, cost of capital, visibility, signaling, and corporate governance considerations, among others (e.g., see the surveys in Karolyi, 1998, 2006). Any of these considerations may explain the imbalance in the number of ADRs in each country in our sample (see column N in Table 1a), as well as in turn be affected by periods of financial turmoil, thus potentially biasing our inference. It is not the objective of this study to examine firms' ADR issuance decisions in emerging economies or whether causes and consequences of financial crises are relevant for those decisions. Nevertheless, when scaling each discrete dependent variable Y_t^m (either E_t^m or P_t^m) by the corresponding total number of ADRs in the sample for country m in week t and estimating Tobit regressions for the resulting ratios, we find the inference below qualitatively unaffected.

[21] Since the weekly currency return volatility σ_{mt}^{FX} for China, Hong Kong, and India (with a mean of 0.89%, 0.03%, and 0.45%, respectively) is generally smaller than for the other countries in the sample, the corresponding estimates of $\widehat{\gamma}_7$ in Table 4 are of greater absolute magnitude.

Table 4
Explaining efficiency breaks E_t^m

	Argentina	Brazil	Chile	China	Hong Kong	India	Indonesia	Israel	Mexico	Peru	Philippines	Poland	Russia	S. Africa	S. Korea	Venezuela
Information																
σ_t^m	−102.5***	−0.2720	−20.65***	11.30***	−124.7**	87.91***	−28.02***	−4.7869	−6.8658*	24.49	−7.7363	8.6179	−24.24**	−17.24***	−37.75***	−111.0***
	(−5.49)	(−0.11)	(−4.69)	(5.22)	(−2.15)	(5.55)	(−4.08)	(−0.36)	(−1.93)	(0.88)	(−0.60)	(0.05)	(−3.36)	(−5.09)	(−3.99)	(−4.52)
σ_t^{US}	138.7***	40.21***	56.54***	17.65*	29.02	−258.5***	85.44***	44.33***	47.32***	128.9***	−41.14**	14.36	83.31***	17.49***	97.61***	212.8***
	(8.26)	(5.78)	(8.55)	(1.92)	(0.51)	(−3.51)	(6.18)	(3.79)	(9.39)	(3.81)	(−2.05)	(0.16)	(3.89)	(4.48)	(8.20)	(4.03)
Trend																
d_t^{m+}	0.0581	−0.0497	0.0115	−0.0638	−0.3400	0.4674*	0.0666	−0.0405	0.0294	−0.0605	−0.1171	0.0741	0.0251	0.0274	−0.0229	−0.1753
	(0.42)	(−0.86)	(0.22)	(−0.98)	(−0.84)	(1.69)	(−0.49)	(−0.44)	(0.61)	(−0.28)	(−0.70)	(0.17)	(0.22)	(0.54)	(−0.21)	(−0.49)
d_t^{m-}	0.2523*	0.0280	0.0319	−0.0287	−0.0330	−0.6019**	0.0254	0.0400	0.0678	−0.1066	−0.0377	−0.0461	−0.0644	−0.0032	0.0438	−0.0830
	(1.85)	(0.38)	(0.59)	(−0.45)	(−0.10)	(−2.44)	(0.19)	(0.37)	(1.19)	(−0.58)	(−0.28)	(−0.13)	(−0.47)	(−0.05)	(0.37)	(−0.35)
d_t^{US+}	−0.0714	−0.0112	−0.0193	0.0066	−0.2809	0.0938	−0.0830	0.0078	−0.0296	−0.0536	0.0462	−0.2597	−0.0862	−0.0169	−0.0816	−0.0736
	(−0.45)	(−0.20)	(−0.38)	(0.11)	(−0.71)	(0.33)	(−0.64)	(0.08)	(−0.65)	(−0.28)	(0.35)	(−0.39)	(−0.67)	(−0.32)	(−0.80)	(−0.22)
d_t^{US-}	0.1240	−0.0003	0.0192	−0.0210	−0.0487	0.3998*	0.0943	0.0030	0.0276	0.0746	−0.0498	0.0297	0.1320	0.0371	0.0804	0.3198
	(1.03)	(−0.01)	(0.30)	(−0.30)	(−0.14)	(1.80)	(0.65)	(0.03)	(0.51)	(0.36)	(−0.28)	(0.08)	(0.99)	(0.55)	(0.75)	(1.28)
Currency and economic environment																
σ_{mt}^{FX}	84.69***	2.5500*	−2.8185	−626.0***	−16,436***	−1,107.8***	−1.4116	−44.79*	10.47***	0.9523	1.7945	−139.8	9.5619***	8.2074***	39.23***	−9.9911
	(6.33)	(1.65)	(−0.41)	(−4.69)	(−3.86)	(−3.79)	(−0.04)	(−1.93)	(3.73)	(0.02)	(0.15)	(−0.29)	(4.15)	(3.25)	(5.24)	(−0.47)
d_{mt}^{FX+}	−0.1407	−0.0446	0.0227	−0.1044	−0.1631	0.5823**	−0.0494	−0.0070	0.0439	0.1129	−0.0884	0.0156	−0.1681	−0.0231	0.1510	0.0556
	(−0.83)	(−0.92)	(0.49)	(−1.27)	(−0.41)	(2.18)	(−0.39)	(−0.07)	(0.89)	(0.54)	(−0.65)	(0.03)	(−1.60)	(−0.44)	(1.45)	(0.26)
d_{mt}^{FX-}	0.0172	0.0189	0.0161	0.0531	0.0208	−0.1677	−0.0097	−0.0049	0.0108	−0.0296	0.0049	−0.0496	−0.2572*	0.0306	0.0643	−0.1265
	(0.13)	(0.26)	(0.26)	(0.57)	(0.05)	(−0.70)	(−0.06)	(−0.05)	(0.20)	(−0.12)	(0.03)	(−0.12)	(−1.72)	(0.51)	(0.64)	(−0.36)
T_t^m	−6.7281**	4.9429***	−6.3495***	n.a.	−6.3460***	−42.74*	2.8350***	−0.6906	−0.1151	−12.15	4.6542***	15,872	1.4961	−1.1511**	2.3423***	−1.8429
	(−2.23)	(4.28)	(−6.57)		(−2.82)	(−1.73)	(6.64)	(−0.70)	(−0.30)	(−0.95)	(4.10)	(0.00)	(1.09)	(−2.29)	(3.26)	(−0.24)
C_t^m	−0.1140***	0.0637***	0.2884***	0.0796***	1.2077***	0.7867	−0.0061	0.0389	0.1432***	0.0248	−0.1874*	−529.6	0.0509**	0.0777***	−0.0012	−0.0050
	(−5.17)	(3.94)	(10.99)	(4.20)	(3.40)	(1.51)	(−0.43)	(1.03)	(6.27)	(0.08)	(−1.72)	(0.00)	(2.23)	(5.43)	(−0.04)	(−0.03)
Investors' activity																
F_t^m	−0.0006	−0.0001	−0.0001	0.0002	−0.00004	−0.0032*	−0.0005	−0.0001	−0.00001	−0.0011	0.0015	0.0036	0.0014	0.0001	−0.0002	0.0012
	(−1.64)	(−1.15)	(−0.43)	(0.85)	(−0.40)	(−1.66)	(−0.49)	(−0.49)	(−0.22)	(−0.81)	(0.92)	(0.21)	(1.11)	(0.39)	(−1.42)	(0.52)
Market liquidity																
L_t^m	−0.1180**	0.0072	0.0879	−0.0009***	−0.1580	−30.55***	11.83	n.a.	−0.0846	0.0640	−0.1671	0.4861	0.0002	0.3019	1.5076	−0.5078
	(−2.09)	(0.25)	(0.07)	(−2.78)	(−0.42)	(−4.30)	(0.94)		(−1.58)	(0.72)	(−1.03)	(0.21)	(0.29)	(1.46)	(0.11)	(−0.08)
L_t^{US}	0.1255	0.2249**	0.2747**	0.2505**	0.3525	0.3368	0.4947*	0.0876	0.4541**	0.5859	0.9115***	0.2167	0.5227**	0.5182**	0.5955**	0.5628
	(0.71)	(2.07)	(2.54)	(1.99)	(0.65)	(0.93)	(1.77)	(0.45)	(5.71)	(1.51)	(3.07)	(0.34)	(2.31)	(4.23)	(3.01)	(1.44)
Expected returns																
D_t^m	0.3828***	−0.0693***	−0.1251***	−0.1627**	0.9551**	0.4890	0.1771*	n.a.	−0.1228**	0.5316	−0.0113	−0.5609	−0.0662	0.1112***	0.0621	−0.0421
	(9.02)	(−4.24)	(−3.76)	(−5.89)	(1.97)	(1.20)	(1.68)		(−2.14)	(1.53)	(−0.06)	(−0.33)	(−0.95)	(4.08)	(0.77)	(−0.43)
D_t^{US}	0.1043	−0.7208***	−0.4806***	0.1752	−3.9257***	4.6872***	−2.5391***	−0.845	−0.8269***	−1.6017***	−0.7179***	0.0750	−1.5681***	−0.1552***	−1.9696***	−1.4139
	(0.31)	(−8.10)	(−7.43)	(0.80)	(−2.53)	(4.03)	(−6.94)	(−0.71)	(−9.21)	(−3.11)	(−3.78)	(0.03)	(−5.24)	(−4.30)	(−5.14)	(−1.24)
t_0	12/7/93	3/15/94	11/28/89	12/21/93	11/19/96	7/14/98	3/21/95	12/1/92	5/31/88	10/15/96	2/9/88	12/5/00	6/3/97	5/27/80	3/14/95	8/17/93
N	487	473	697	485	333	247	420	534	775	338	791	122	305	1,210	421	503
R_P^2	95.86%	94.55%	95.95%	88.53%	90.71%	77.73%	81.98%	85.66%	97.68%	59.76%	67.65%	98.48%	63.08%	88.01%	75.10%	93.04%
R_D^2	94.45%	92.81%	96.09%	91.41%	88.79%	78.91%	85.67%	82.74%	97.39%	71.13%	75.03%	98.35%	70.78%	85.21%	82.93%	90.32%
χ^2	132.5***	74.87***	68.79***	146.9***	21.65	67.17***	53.40***	28.25***	102.8***	39.04***	119.0***	0.8687	71.90***	92.67***	104.4***	38.57***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.1172)	(0.0000)	(0.0000)	(0.0083)	(0.0000)	(0.0006)	(0.0000)	(1.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0007)

This table reports maximum likelihood estimates for the Poisson regression model of Eqs. (4) and (5) when $Y_t^m = E_t^m$ for country m over an interval starting on the earliest potential break date (t_0). N is the resulting number of observations. R_P^2 is the Pearson R^2; R_D^2 is the Deviance R^2; χ^2 is a goodness-of-fit statistic; its p-values are in parentheses. A " * ", " ** ", or " *** " indicate significance at the 10%, 5%, or 1% level, respectively.

Table 5
Explaining pricing breaks P_t^m

	Argentina	Brazil	Chile	China	Hong Kong	Hungary	India	Indonesia	Israel	Mexico	Philippines	Russia	S. Africa	S. Korea	Venezuela
Information															
σ_t^m	−63.03***	−24.58***	−7.6265	−2.4965	−2.7346	10.35	30.67**	−27.20***	−35.46*	−16.71***	−12.50	−14.48*	4.0104	−25.20***	−213.0***
	(−4.79)	(−5.72)	(−1.61)	(−1.14)	(−0.25)	(0.56)	(2.56)	(−4.02)	(−1.91)	(−4.32)	(−0.95)	(−1.84)	(1.35)	(−3.26)	(−3.21)
σ_t^{US}	196.2***	114.0***	75.96***	50.70***	51.25*	117.6***	72.11*	78.57***	48.29***	25.54***	−36.88*	72.14***	−0.0310	19.45	141.5**
	(12.63)	(13.36)	(10.90)	(8.05)	(1.95)	(2.97)	(1.81)	(5.81)	(3.56)	(4.84)	(−1.80)	(3.01)	(−0.01)	(1.46)	(2.08)
Trend															
d_t^{m+}	0.0601	0.0700	0.0121	−0.0913	−0.1270	0.1449	0.0363	−0.0523	−0.0179	0.0286	−0.0771	0.0392	0.0073	0.0462	−0.5077
	(0.48)	(1.14)	(0.22)	(−1.31)	(−0.88)	(0.66)	(0.21)	(−0.39)	(−0.17)	(0.61)	(−0.46)	(0.31)	(0.15)	(0.38)	(−0.85)
d_t^{m-}	0.1821*	0.1122	0.0479	−0.0030	−0.0153	0.0367	−0.1596	0.0023	0.0130	0.0332	−0.0486	0.0318	0.0125	−0.0412	−0.2263
	(1.75)	(1.52)	(0.84)	(−0.04)	(−0.12)	(0.19)	(−0.94)	(0.02)	(0.10)	(0.60)	(−0.36)	(0.22)	(0.22)	(−0.32)	(−0.59)
d_t^{US+}	−0.1369	−0.0941	−0.0328	−0.0074	0.0047	−0.2441	−0.0384	−0.0486	−0.0708	−0.0397	0.0166	−0.0408	0.0041	−0.0497	0.5241
	(−1.08)	(−1.51)	(−0.60)	(−0.12)	(0.04)	(−1.08)	(−0.21)	(−0.39)	(−0.68)	(−0.93)	(0.22)	(−0.30)	(0.09)	(−0.48)	(0.94)
d_t^{US-}	0.1609	0.0082	0.0491	0.0204	0.0019	0.0966	0.1481	0.0861	−0.0104	0.0334	−0.0476	0.0803	0.0187	0.0149	0.3339
	(1.62)	(0.13)	(0.77)	(0.28)	(0.01)	(0.44)	(0.88)	(0.60)	(−0.08)	(0.63)	(−0.26)	(0.52)	(0.28)	(0.11)	(0.90)
Currency and economic environment															
σ_{mt}^{FX}	39.68***	7.7168***	−33.21***	−14.85***	95.22	94.51	−139.3	−2.9209	10.37	21.99***	4.3217	7.7446***	−0.0811	21.22***	20.04
	(4.16)	(4.66)	(−4.41)	(−2.74)	(0.07)	(0.83)	(−1.61)	(−0.09)	(0.36)	(8.63)	(0.35)	(3.19)	(−0.03)	(3.50)	(0.68)
d_{mt}^{FX+}	−0.1179	−0.0334	0.0123	−0.0586	0.0287	−0.0340	0.0916	−0.0218	−0.0444	0.0143	−0.0792	−0.1592	−0.0102	0.0539	−0.1665
	(−0.71)	(−0.64)	(0.25)	(−0.67)	(0.24)	(−0.16)	(0.57)	(−0.17)	(−0.38)	(0.30)	(−0.58)	(−1.33)	(−0.21)	(0.49)	(−0.50)
d_{mt}^{FX-}	−0.1789	0.0447	0.0166	0.0228	0.0449	−0.0510	−0.1356	−0.0282	0.0957	0.0048	0.0074	−0.2250	−0.0097	0.0202	−0.1215
	(−1.40)	(0.62)	(0.26)	(0.24)	(0.25)	(−0.22)	(−0.65)	(−0.16)	−6.9352***	1.0543***	4.6334***	0.4027	−1.8805***	1.4954*	4.9430
T_t^m	−0.1608	−3.3730***	−2.9894***	n.a.	−1.1240**	−1.3260	−3.1034	2.8525***	(−6.04)	(2.84)	(4.01)	(0.27)	(−3.47)	(1.95)	(0.37)
	(−0.08)	(−2.55)	(−2.97)		(−2.21)	(−0.96)	(−0.35)	(6.93)	0.3432***	0.1780***	−0.1740	0.0172	0.1716***	0.0752***	0.2422
C_t^m	−0.0410**	0.1786***	0.1555***	0.0851***	0.1521**	−0.0242	−0.1724	−0.0041	(6.54)	(7.31)	(−1.58)	(0.69)	(11.18)	(2.76)	(0.99)
	(−2.23)	(7.16)	(5.56)	(4.46)	(2.11)	(−0.14)	(−0.78)	(−0.29)							
Investors' activity															
F_t^m	−0.0003	−0.0001	−0.0003	−0.0002	−0.00001	−0.0013	0.0016	−0.0007	0.00003	0.00004	0.0014	0.0008	0.0003	0.00003	−0.0010
	(−1.10)	(−1.28)	(−1.06)	(−0.77)	(−0.19)	(−0.52)	(1.55)	(−0.67)	(0.17)	(0.54)	(0.87)	(0.56)	(1.20)	(0.22)	(−0.21)
Market liquidity															
L_t^m	−0.1463***	0.0226	0.9328	−0.0004	0.0436	−54.75	−0.3121	13.01	n.a.	−0.4897***	−0.1484	0.0001	−0.0360	−3.3398	−9.7260
	(−2.95)	(0.82)	(0.78)	(−1.29)	(0.59)	(−1.56)	(−0.10)	(1.05)		(−5.95)	(−0.91)	(0.12)	(−0.19)	(−0.18)	(−0.57)
L_t^{US}	0.7352***	0.1014	0.3185***	0.3137***	0.1944	0.4797	0.0901	0.4589*	0.840**	0.5387***	0.8209***	0.4865*	0.4605***	0.2083	0.4482
	(4.62)	(0.91)	(2.94)	(2.63)	(0.75)	(1.11)	(0.28)	(1.66)	(2.23)	(6.78)	(2.73)	(1.85)	(3.74)	(0.88)	(0.81)
Expected returns															
D_t^m	0.1487***	0.1154***	−0.2667***	0.0155	0.1383	−0.2260	−0.6431***	0.1870*	n.a.	−0.4309***	−0.0195	−0.1540*	−0.0881**	0.0299	0.0697
	(4.84)	(6.98)	(−6.11)	(0.62)	(1.64)	(−0.88)	(−2.59)	(1.80)		(−6.08)	(−0.10)	(−1.93)	(−2.50)	(0.33)	(0.56)
D_t^{US}	−2.4326***	−1.7685***	−0.2369***	−0.9133***	−1.0599***	−1.9280***	1.2889	−2.4323***	−0.9282***	−0.6964***	−0.8153***	−1.0930***	−0.3071***	−1.1563***	−2.4646
	(−9.19)	(13.92)	(−3.55)	(−7.61)	(−2.73)	(−2.63)	(1.51)	(−6.96)	(−6.20)	(−7.14)	(−4.05)	(−3.29)	(−8.35)	(−4.46)	(−1.35)
t_0	12/7/93	3/15/94	11/28/89	12/21/93	11/19/96	4/14/98	7/14/98	3/21/95	12/1/92	5/31/88	2/9/88	6/3/97	5/27/80	3/14/95	8/17/93
N	487	473	697	485	333	260	247	420	534	775	791	305	1,210	421	503
R_P^2	94.24%	95.16%	95.23%	90.23%	32.92%	41.91%	63.04%	81.68%	72.24%	90.29%	68.87%	53.57%	83.76%	69.08%	92.37%
R_D^2	92.17%	95.58%	94.85%	91.40%	51.59%	57.70%	70.03%	85.64%	78.51%	91.03%	76.09%	66.83%	82.93%	70.88%	88.50%
χ^2	140.2***	102.7***	92.52***	95.83***	38.63***	27.88**	37.28***	52.79***	105.3***	447.5***	117.1***	59.43***	207.0***	19.79	33.27***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0007)	(0.0223)	(0.0012)	(0.0000)	(0.0000)	(0.0006)	(0.0000)	(0.0003)	(0.0000)	(0.1802)	(0.0043)

This table reports maximum likelihood estimates for the Poisson regression model of Eqs. (4) and (5) when $Y_t^m = P_t^m$ for country m over an interval starting on the earliest potential break date (t_0). N is the resulting number of observations. R_P^2 is the Pearson R^2; R_D^2 is the Deviance R^2; χ^2 is a goodness-of-fit statistic; its p-values are in parentheses. A " * ", " ** ", or " *** " indicate significance at the 10%, 5%, or 1% level, respectively.

exchange rate on the efficiency of emerging ADR markets is symmetrically unrelated to its prolonged unidirectional swings: Coefficients for d_{mt}^{FX+} and d_{mt}^{FX-} in Table 4 are statistically insignificant for most of the countries in the sample.

As discussed in Section 5, Tables 2 and 3 report that many of the statistically significant breaks in Eq. (1) are associated with statistically significant breaks in Eq. (2) in proximity of financial crises, especially when these breaks entail a deterioration in the efficiency and pricing of the corresponding ADRs. Consistently, Table 5 indicates that, in those circumstances, the likelihood of ADR prices to become more (less) sensitive to local (global) sources of risk is similarly related to both domestic and U.S. stock return volatility and the dynamics of the domestic currency: $\widehat{\gamma}_1<0$, $\widehat{\gamma}_2>0$, $\gamma_7>0$, $\widehat{\gamma}_8\approx 0$, and $\widehat{\gamma}_9\approx 0$, respectively, in Eq. (5) for P_t^m. Not only the occurrence but also the interaction between efficiency and pricing breaks may be related to the trading activity of investors and speculators during periods of financial turmoil. For instance, much trading in emerging market equities already takes place on the NYSE, NASDAQ, or AMEX during "tranquil" times (e.g., Hargis, 1997). More trading, especially if informed, may shift toward these and other more liquid venues during periods of turmoil, ultimately affecting both emerging ADRs' arbitrage relation (Eq. (1)) and the covariance of their returns with local and U.S. equity returns (Eq. (2)). Accordingly, we find that greater divergence in liquidity (lower Lm and/or higher LUS) is often associated with both efficiency and pricing breaks — i.e., $\widehat{\gamma}_{13}<0$ and/or $\widehat{\gamma}_{14}>0$ in Tables 4 and 5 — even when the former are in the direction of the law of one price (in Brazil, China, Philippines, and Russia, but not India).[22]

Somehow surprisingly, capital flight — measured by net monthly equity capital flows (F_t^m) — does not impact the clustering of arbitrage and pricing breaks in emerging ADR markets, with the sole exception of India. Furthermore, most of the coefficients for the upward and downward trend dummies in the domestic and U.S. stock markets defined in Section 6 — $\widehat{\gamma}_3$ to $\widehat{\gamma}_6$ in Eq. (5) — are statistically insignificant. Nonetheless, these variables may not adequately capture the extent to which capital controls and financial constraints are curtailing the ability of domestic and foreign investors to fully arbitrage price wedges between ADRs and their underlying stocks at each point in time. For instance, although virtually all of the countries in our sample gradually relaxed capital mobility and foreign ownership restrictions during the 1990s (e.g., Bekaert et al., 2002), some of their governments also abruptly reintroduced many of these restrictions in proximity of financial crises, most notably in Argentina between 2001 and 2002.[23] Yet again, when considering more direct measures of the time-varying intensity of country-level capital controls and local market openness — such as the Edison and Warnock (2003) index and the Chinn and Ito (2006) index (as in Auguste et al., 2006) — we find weak or no evidence (not reported here) that structural shifts in ADR efficiency and pricing relationships are more likely when the activity of investors and arbitrageurs is more hindered.[24]

Despite these results, shifts in investors' expectations do appear to play an important role for the dynamics of E_t^m and P_t^m: Lower expected returns in the U.S. (D_t^{US}) and in the domestic market (D_t^m) almost uniformly increase the likelihood of the law of one price to be violated and of a significant $\Delta\widehat{b}_i^{US}<0$ to be observed. ADR markets are instead on average more efficient, but more sensitive to local sources of risks, in response to higher D_t^m. Lastly, we find that a shrinking (expanding) trade sector — i.e., lower (higher) T_t^m — also leads often to the clustering of negative (positive) shifts in $\widehat{b}_i$ and $\widehat{b}_i^{US}$, and/or positive (negative) shifts in $\widehat{b}_i^m$ — i.e., $\widehat{\gamma}_{10}<0$ ($\widehat{\gamma}_{10}>0$) in Eq. (5), especially in Latin America and southeast Asia. Yet, deterioration (improvement) in the efficiency and pricing of ADRs are generally less (more) likely to occur during worsening (better) macroeconomic conditions (as proxied by declining S&P sovereign ratings C_t^m): $\widehat{\gamma}_{11}>0$ in Tables 4 and 5 for most countries. These results indicate that the healthier economic environment of an emerging country and the ensuing attention from the financial community ameliorate the efficiency and pricing of its ADR market only if accompanied by greater trade linkages with the rest of the world.[25]

Overall, the above evidence — subject to the caveats discussed earlier and some country-level heterogeneity — suggests that many of the factors usually associated with the occurrence of financial crises in emerging economies, in particular increasing uncertainty and disagreement among investors, greater exchange rate volatility, dwindling trade balances, unfavorable expectations, and worsening liquidity (but not stock market trends, currency devaluations, or flight of foreign capital) also contribute to make their emerging equity markets less efficient and — assuming the validity of the two-factor pricing model of Eq. (2) — more de facto segmented.

[22] Similar inference can be drawn when using the average proportion of zero return weeks for the underlying local stocks of the ADRs in our sample as a proxy for those stocks' degree of liquidity (as in Lesmond, 2005).

[23] Nevertheless, the majority of estimated regime shifts in Eqs. (1) and (2) for Argentina took place *before* new restrictions to capital flight (described in Levy Yeyati et al., 2004; Auguste et al., 2006) were announced and implemented. Our evidence of lower efficiency and segmentation of the Argentine ADR market in 2001 (in Tables 1–3 and Fig. 1) is consistent with the argument in Melvin (2003) that domestic investors moved funds out of Argentina, despite stringent capital controls, by purchasing shares of local companies cross-listed in the U.S. and then converting them in ADRs, until such conversions were permitted (March 25, 2002).

[24] In unreported analysis, we also consider whether, in the presence of less-than-perfect legal substitutability between ADRs and their associated shares, informed risk arbitrage in emerging markets may become less attractive to arbitrageurs (e.g., Morck et al., 2000; Aggarwal et al., 2007), thus affecting these markets' efficiency and pricing. We find no relation between country-level and firm-level measures of investor protection and corporate governance (such as those developed by La Porta et al., 1998, and by Credit Lyonnais Securities Asia, 2001) and the likelihood and extent of the estimated efficiency and pricing breaks in Tables 1a and 2a.

[25] The likelihood of an emerging country's ADRs to experience persistent regime shifts in their arbitrage and pricing relationships may also be explained by financial contagion – i.e., non-fundamental cross-country spillover – due to the trading activity of investors and speculators (for a review see Kaminsky et al., 2003; Kallberg et al., 2005). We test for this possibility by analyzing whether the number of emerging markets in our sample simultaneously experiencing more efficiency or pricing breaks than justified by the fundamental model of Eqs. (4) and (5) is greater than if due to chance, in the spirit of Kallberg et al. (2005). The resulting test statistics (available on request) overwhelmingly reject the alternative hypothesis of financial contagion, except for pricing breaks during the Mexican peso crisis.

7. Conclusions

Financial crises have received increasing attention from the economic and financial literature, not least in light of their tremendous social and allocational costs. Determining the nature of financial crises is crucial to explaining their occurrence and formulating policy recommendations for their prevention (e.g., Schwartz, 1986; Mishkin, 1992; IMF, 1998; Bordo et al., 2000). The main objective of this paper was to contribute to the understanding of the anatomy of these crises by conducting an empirical investigation of the impact of their occurrence on the efficiency and pricing of the emerging ADR market.

Specifically, we first identified two fundamental relationships that arbitrage and capital mobility considerations impose on returns of depositary receipts and their local counterparts. The first one — between the dollar return for an ADR and the dollar return for the underlying security — arises from the law of one price, ceteris paribus for various financial frictions, and should always hold unless "normal" market conditions are violated. The second one — between ADR returns and local and global sources of risk — is the result of perfectly open and fully integrated financial markets, conditional upon the validity of a standard two-factor ADR pricing model. We then tested for a single, persistent regime shift in each of these relations in our sample of ADRs from issuers in 20 emerging countries using the non-parametric technique of Bai et al. (1998).

We provided evidence that during recent episodes of financial turmoil — Mexico (1994), East Asia (1997/1998), Russia (1998), Brazil (1999), Turkey (2001), and Argentina (2001/2002) — those normal market conditions were often violated. In particular, we found that a majority of the depositary receipts in our sample experienced statistically and economically significant structural shifts in their efficiency and pricing relationships over the past decade or so. The estimated properties and chronology of these breaks suggest that financial crises involving emerging economies were accompanied by lower efficiency and greater segmentation of their ADR markets.

Finally, using several regression models, we offered support for some of the available explanations of financial crises in emerging economies by showing that the estimated regime shifts in each country in our sample were more likely when uncertainty among investors was greater and both financial and macroeconomic conditions were deteriorating. However, those efficiency and pricing breaks were also generally unrelated to capital controls, prolonged stock market runs, currency depreciations, and foreign capital flight.

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