Dating the integration of world equity markets ☆

Geert Bekaert[a,d], Campbell R. Harvey[b,d,*],
Robin L. Lumsdaine[c,d]

[a] *Columbia University, New York, NY 10027, USA*
[b] *Duke University, Durham, NC 27708, USA*
[c] *Deutsche Bank AG, London, EC2N 2DB, UK*
[d] *National Bureau of Economic Research, Cambridge, MA 02138, USA*

Regulatory changes that appear comprehensive will have little impact on the functioning of a developing market if they fail to lead to foreign portfolio inflows. We specify a reduced-form model for a number of financial time series and search for a common, endogenous break in the data generating process. We also estimate a confidence interval for the break. Our endogenous break dates are accurately estimated but do not always correspond closely to dates of official capital market reforms. Indeed, the endogenous dates are usually later than official dates, highlighting the important distinction between market liberalization and market integration.

☆ We have benefited from discussions with and the comments of Martijn Cremers, Vihang Errunza, Lars Hansen, Andrew Harvey, Ming Huang, Luis Viceira, and seminar participants at Brown, Cambridge, Chicago, Darden, Erasmus, Maryland, Indiana, and New York University, as well as participants at the American Finance Association Meetings, the European Finance Association Meetings in Barcelona, and the Chazen Institute Conference on International Valuation at Columbia University. The paper has especially benefited from the detailed comments of the referee. We wish to thank Eric Engstrom, Andrew Roper, and Diego Valderrama, our RAs, for their help. This research was initiated while Lumsdaine was a National Fellow at the Hoover Institution. The opinions expressed are the authors' and do not represent those of Deutsche Bank or its affiliates.

*Corresponding author. Tel. +1-919-660-7768; fax: +1-919-660-8030.
E-mail address: cam.harvey@duke.edu (C.R. Harvey).

1. Introduction

In financially integrated markets, domestic investors are able to invest in foreign assets and foreign investors in domestic assets; hence, assets of identical risk command the same expected return, regardless of trading location. Moving from a segmented regime to an integrated regime affects expected returns, volatilities, and correlations with world factors, all of which are important for both risk analysis and portfolio construction. Consequently, the concept of market integration is central to the international finance literature.

Market integration also plays a key role in international and development economics. International economists, however, focus on the potential welfare gains of market integration, in terms of risk sharing benefits (see Cole and Obstfeld, 1992; Lewis, 1996; van Wincoop, 1994). In the development economics literature, Obstfeld (1994), Bekaert et al. (2001a,b), and Henry (2000a) have started to trace the investment and growth benefits of financial market integration. The interplay between the financial sector and growth is examined in Devereux and Smith (1994) and Levine and Zervos (1996), among many others.

With the opening of so many emerging markets in the last decade, history now offers a unique experiment to explore the economic and financial effects of market integration. Not surprisingly, a literature has developed attempting to measure the macroeconomic and financial effects of market integration (see Bekaert and Harvey, 1995, 1997, 1998, 2000; Aggarwal et al. 1999; De Santis and İmrohoroğlu, 1997; Richards, 1996; Levine and Zervos, 1996; Kim and Singal, 2000; Henry, 2000a,b; Domowitz et al. 1997).[1]

But how do we measure the process of market integration? Indeed, how do we test the equilibrium models of risk sharing? How do we measure the growth effects of market integration? A prerequisite to these questions is the date that a market becomes integrated. The dating question is the subject of our research.

The dating of market integration is difficult. The capital market liberalization process is a complex process and it is unlikely that "dates" of capital market reforms will correspond to the true date of market integration. For example, there are often ways to circumvent capital controls. Investors can access markets indirectly through American Depositary Receipts (ADRs) or country funds, even though the market is technically closed to foreign investors. Liberalization may occur in stages and be a gradual process. Finally, some policy changes may be anticipated well in advance while others lack credibility.

There are potential solutions to these problems. One option is to specify a tightly parameterized model of the process of dynamic integration. For example, Bekaert and Harvey (1995) use a regime-switching framework to model gradual changes in market integration. However, these models are difficult to specify and are often statistically rejected. In addition, international asset pricing models typically fail to

[1] There is an older theoretical and empirical literature on market integration going back to Solnik (1974), Stehle (1977), Stulz (1981), Errunza and Losq (1985), Eun and Janakiramanan (1986), Jorion and Schwartz (1986), and Errunza et al. (1992).

match the home asset preference that investors display, even when markets are seemingly perfectly integrated.[2]

In this paper, we offer an alternative approach that results in market liberalization dates, with confidence intervals, for 20 countries. Our methodology is new in two respects. First, in contrast to Bekaert and Harvey's (1995) focus on returns, we look at a series of financial and macroeconomic variables that are likely related to the integration process. For example, we consider net equity capital flows as well as variables such as dividend yields that may capture the permanent price effects that market integration entails. Indeed, returns are noisy in emerging markets and it is important to expand the scope of examination to other variables.

Second, we do not take a stand on an asset pricing model, but simply assume that the variables before and after market integration follow a stationary process that is well described by a vector autoregression (VAR). Our methodology then exploits the new technique of Bai et al. (1998) to find endogenous break points for the VAR parameters. Since the model is in reduced form, it is important to let all parameters change. The methodology also yields a break date with a confidence interval. Interestingly, Bai et al. show that the confidence intervals around break dates can be tightened considerably by adding different series that break at the same time. One of the unique aspects of our research is that we simultaneously examine a number of economic time series in our search for structural breaks.

The paper is organized as follows. In the second section, we describe the liberalization process and formulate formal hypotheses. The third section details the econometric methodology and examines the finite sample properties of the structural break identification procedure. The data are described in the fourth section. The fifth section contains our empirical results. We present both univariate and multivariate break points for the time series that we examine. The sixth section provides an economic interpretation of the statistical break points and discusses the effects of financial market integration on a wide range of financial and economic indicators. Some concluding comments and caveats are offered in the final section.

2. The capital market liberalization process

2.1. The impact of liberalizations

The liberalization process is extremely complex and there is no established economic model that guides us. That is, while there are general equilibrium models of economies in integrated states and segmented states, there is no model that specifies the economic mechanism that moves a country from segmented to integrated status.

In the standard mean-variance integration/segmentation model (see Stapleton and Subrahmanyam, 1977; Errunza and Losq, 1985; Alexander et al. 1987), prices under integration are decreasing in the covariance between world and local cash flows

[2] See, for example, Tesar and Werner (1995) and Kang and Stulz (1997).

whereas in a segmented world, only local cash flow volatility matters for pricing. Likewise, in an integrated world, the expected returns are linked to the covariance with world market returns as opposed to the covariance with local return volatility under segmentation (see Bekaert and Harvey, 1995). Because the volatility of emerging market returns is much higher than the covariances with world market returns, these models suggest that prices will jump up on the announcement of a liberalization and expected returns will decrease. The size of the jump should be related to both the credibility of the government's announcement (and its policies in general) and the diversification benefits to be gained from integrating the market. When the market finally liberalizes, foreign capital flows in and the price rises again since all uncertainty is resolved. This last price rise can be small if the announcement was credible. It is possible that when a market is opened to international investors, it will become more sensitive to world events (covariances with the world will increase). Even with this effect, it is likely that these covariances are still much smaller than the local variance, which would imply rising prices. Indeed, Bekaert and Harvey (2000), Henry (2000a), and Kim and Singal (2000) empirically document these returns to integration and implications for the cost of capital.

This simple intuition suggests that we will see breaks not only in returns but maybe even more clearly in variables such as dividend yields and ratios of market capitalization to GDP that reflect permanent price changes. Of course, the underlying models ignore many interesting dynamic effects. As a result, we concentrate our discussion on a large set of factors that are likely to be consistent with integration. It is possible that none of these factors when examined individually will reveal market integration. However, considered as a group—or in subgroups—they may be useful in determining the degree of integration.

For example, the liberalization process might be reflected in local market activity. As foreigners are allowed to access the local market, liquidity can increase along with trading volume. We may be able to directly observe the interest of foreign investors by examining the net equity capital flows to these markets.

There could also be some structural changes in the market. For example, if the cost of capital decreases, new firms might present initial public offerings. The stock market concentration will decrease as a result of these new entrants. In addition, individual stocks might become less sensitive to local information and more sensitive to world events. This can cause the cross-correlation of the individual stocks within a market to change.

The liberalization process is intricately linked with the macroeconomy. Liberalization of financial markets could coincide with other economic policies directed at inflation, exchange rates, and the trade sector (see Henry, 2000a for details). International bankers who establish country risk ratings likely view liberalization as a positive step. Hence, these ratings will contain valuable information regarding the integration process as well as the credibility of reforms.

As mentioned earlier, examining any single or even a small group of these indicators can be misleading. For example, capital flows might not reveal a substantial change because foreign investors are able to access the domestic market via ADRs and country funds. However, it makes sense that a comprehensive examination of a number of indicators will provide additional information.

There are of course important timing issues to be considered.[3] Market prices could change upon announcement of the liberalization or as soon as investors anticipate a potential future liberalization. However, foreign ownership can only be established when allowed by the authorities. Although we look at breaks in all of our variables, we are careful in grouping them, taking into account such timing issues in the interpretation of our results.

2.2. Economic events and liberalization

Bekaert and Harvey (1998) provide a detailed examination of the economic events that could impact the liberalization process for each of the 20 countries in our sample and present a chronology of both country fund launchings and ADR issues. We use this information to give economic content to our statistical tests.

As an example, consider the information collected for Colombia. In early 1991, a series of regulatory moves was directed at easing the ability of foreign-owned corporations to repatriate profits. There were significant economic reforms in mid-1991 that reduced tariffs in two steps. Much of the economy's external debt was refinanced in March 1991.

In October 1991, two important economic events occurred. The first involved the deregulation of the peso. The currency was allowed to float freely and exchange controls were eased. At the same time, Resolution 51 allowed foreigners to invest up to 10% in the local equity market. In November 1991, Resolution 52 eliminated the 10% cap on equity investment. By December 1991, the telecom industry was privatized.

In April 1992, the Colombian Investment Company (a country fund) was listed in Luxembourg. In February 1993, the first ADR was offered to institutional investors as a rule 144A private placement (announced in December 1992). In September 1993, a number of currency-related reforms were passed: the exchange controls that required that pesos be converted to U.S. dollars and then translated into other currencies were eliminated; local firms were given the ability to obtain short-term loans abroad; local firms could now make peso loans to foreign-owned corporations based in Colombia; the U.S. dollar could be used for domestic transactions; and the use of derivative securities on currencies was approved. In November 1994, an ADR of a Colombian company was listed on the NYSE.

What can structural break tests teach us about a process as complex as the capital market liberalization process in Colombia just described? First, they can reject the null hypothesis of no structural break. That is important in itself. Indeed, there is no guarantee that the liberalizations that took place effectively integrated this emerging market into world capital markets. If we fail to reject the null of no structural break

[3] For example, Foerster and Karolyi (1999), examining ADRs, find share price reactions that precede the cross-listing date. Miller (1999) details the price impact of announcements of intention to list rather than the actual listings for a large sample of ADRs.

in our stock market variables, we must conclude that the liberalization process that took place over the last decade had little effect on the return-generating process.

Second, when there is a significant break, we can also investigate the break date with the associated confidence interval. Knowing that the return-generating process underwent a break is a necessary but not sufficient condition for concluding that capital market liberalizations actually served to integrate the capital market. We need to examine the timing of the break using the kind of information assembled for every country in Bekaert and Harvey (1998) to determine whether it was caused by a capital market liberalization or another major event such as the privatization of a major company. One might suspect that when the market integration is gradual and occurs in small increments, the confidence interval around the break is large. However, it is also possible that gradual policy changes induce discrete and sudden changes in market prices and other financial variables, since agents anticipate further policy changes. Our methodology allows us to distinguish between these cases on a country-by-country basis.

Third, since we estimate the reduced-form dynamics of the variables before and after the liberalization, we can examine what changes the break induces in the stochastic process governing our variables and investigate whether these changes are consistent with our evidence and theory regarding the effects of market integration.

Finally, as recent events in Southeast Asia have amply shown, the market integration process can be reversed. For example, Malaysia effectively suspended convertibility of its currency in late 1997. This is potentially quite damaging for our methodology. What is the value of our tests of market integration if the process is better described as a regime-switching model (see, e.g., Bekaert and Harvey, 1995)? At this point, we do not have a full answer and this question merits more research. Nevertheless, our intuition is that if the regimes are persistent enough, a regime switch acts as a near-permanent break and our tests will have power to detect when the switch occurred. If there are frequent switches, we suspect our confidence interval around the break will be wide. A serious problem arises if there are exactly two or three switches, in which case our tests may not be that informative. We will come back to this issue in Section 5.4 and discuss some robustness checks. One potential approach is to investigate the tests for multiple breaks in univariate time series, as in Bai and Perron (1998).[4]

3. Econometric methods

Our goal is to test for structural breaks in multiple time series that are potentially impacted by the integration process. The work of Banerjee et al. (1992) and Bai et al. (1998) is ideally suited for this purpose. These papers contain three key observations:

[4] However, to date there are no methods that combine the attractiveness of testing for multiple breaks with the benefits of a multivariate framework; as the latter is the focus of this paper, here we simply plot our Wald tests for the structural break over the whole period and assess the chance that there are multiple breaks informally.

(i) tests can be constructed to determine whether or not a structural break occurred in the data, (ii) the precision with which a potential break date is estimated is a function not of the number of observations in a single series but of the number of series in a multivariate framework that experience the same break date, and (iii) confidence intervals can be computed enabling inference about the break date. They demonstrate this for both stationary vector autoregressions and cointegrated systems. We begin by sketching the intuition for these observations in a simple multivariate model with a known break date and no regressors and generalize to the multivariate case with an unknown break date and stationary regressors as is relevant for our purposes. We end with the results of a Monte Carlo analysis of the behavior of the test statistics in small samples.

3.1. Unknown break date, no regressors

Let y_t be an $n \times 1$ vector following the process $y_t = C + e_t$ and consider estimating the following model:

$$y_t = C_0 + C_1 \mathbf{1}(t > k) + \varepsilon_t, \tag{1}$$

where $\mathbf{1}(A)$ is an indicator function equal to one if event A is true and zero otherwise. This model allows us to test whether a structural break occurred at date k under the alternative; under the null hypothesis there is no break. Let $\hat{\Sigma}$ denote the estimator of the variance–covariance matrix of the $n \times 1$ vector ε_t and $\tau = k/T$ be the fraction of the sample at which the break occurs. Then a test of the null hypothesis that no break occurred in all series at date k can be constructed as a Wald-type test using $\tau(1-\tau)\hat{C}_1'[\mathrm{Var}(\hat{C}_1)]^{-1}\hat{C}_1$. The limiting distribution of this statistic is a $\chi^2(n)$, since there are n elements in C_1.

Suppose now that we want to choose between all possible break dates. To determine the date at which the break is most likely to have occurred, we could compute our $\chi^2(n)$ statistic at each possible date and then designate the maximum over this sequence of statistics as the break date. If this maximum statistic exceeds some critical value, we reject the null hypothesis of no break in favor of a break. To find the correct critical value, we could generate T-tuples of random χ^2 variables (where T is our sample size), take the maximum, sort, and use the value associated with the 95th percentile.

To find the break date, at each possible date k a standard test statistic is computed as in (1), resulting in a sequence of test statistics, which we denote by $F_T(k)$. In their Theorem 1, Bai, Lumsdaine, and Stock (BLS hereafter) apply the functional Central Limit Theorem to this sequence to show that its limiting distribution is

$$F^*(\tau) = \{\tau(1-\tau)\}^{-1}\|W(\tau) - \tau W(1)\|^2, \tag{2}$$

where $\|\cdot\|$ represents the Euclidean norm and $W(\cdot)$ is a vector of n independent standard Brownian motion processes. Because the maximum is a continuous function, it follows from the continuous mapping theorem that $\max_k F_T(k)$ converges in distribution to $Sup\, F^*(\tau)$. This test is called the Sup-Wald test. The observation $\hat{k}$ that maximizes $F_T(k)$ is our estimated break date. While the break date cannot be consistently estimated in this context, the estimate of the fraction of the sample at which the break occurs, $\hat{\tau} = \hat{k}/T$, is a consistent estimator for the true fraction, $\tau_0 = k_0/T$.

BLS provide a table of critical values for the Sup-Wald test derived from a discrete approximation to the continuous distribution given in Eq. (2). BLS further study the joint limiting distribution of $\hat{k}$, $\hat{C}(\hat{k})$, and $\hat{\Sigma}(\hat{k})$, the maximum likelihood estimators of the break date, the vector C, and the variance, Σ (for notational simplicity, we will suppress dependence on $\hat{k}$ by writing $\hat{C}$ and $\hat{\Sigma}$ from now on). They establish that

$$(\hat{C}_1' \hat{\Sigma}^{-1} \hat{C}_1)(\hat{k} - k_0) \Rightarrow V^*, \tag{3}$$

which has a well-defined limiting distribution under the alternative that the true break occurs at k_0. In particular, V^* is distributed as $\text{argmax}_v\ (W(v) - \frac{1}{2}|v|)$, with the limiting density of V^* given in Picard (1985) as

$$\gamma(x) = \tfrac{3}{2}\mathrm{e}^{|x|}\Phi(-\tfrac{3}{2}\sqrt{|x|}) - \tfrac{1}{2}\Phi(-\tfrac{1}{2}\sqrt{|x|}), \tag{4}$$

where $\Phi(\cdot)$ is the cumulative normal distribution function and W is now a one-dimensional two-sided Brownian motion on $(-\infty, \infty)$.[5] Confidence intervals of width $(1-\pi)$ can be constructed using

$$\hat{k} \pm \alpha_{\pi/2}(\hat{C}_1' \hat{\Sigma}^{-1} \hat{C}_1)^{-1}, \tag{5}$$

where $\alpha_{\pi/2}$ is the $(1 - \frac{1}{2}\pi)$th quantile of V^*.

Eq. (5) implies that the confidence interval shrinks with more series that have the same break date. In the simplest case where the elements of ε_t are uncorrelated so that Σ is a diagonal matrix, the confidence interval is

$$\hat{k} \pm \alpha_{\pi/2}\left(\sum_{i=1}^{n} \frac{\hat{C}_{1i}^2}{\hat{\Sigma}_i}\right)^{-1},$$

where $\hat{\Sigma}_i$ is the ith diagonal element of $\hat{\Sigma}$ and $\hat{C}_{1i}$ is the ith element in $\hat{C}_1$. Since each term in the summation is strictly positive, the addition of another equation will only decrease (and never increase) the width of the confidence interval. In particular, when the magnitude of the break (C_{1i}) and variance of ε_{it} (Σ_i) are the same for each series, the confidence interval shrinks at the rate $1/n$. This differs from the usual case of constructing confidence intervals for regression coefficients that are the same across equations in which confidence intervals shrink at rate $1/\sqrt{n}$. In addition, even if some of the series do not contain a break, so that $C_{1i} = 0$ for some equations, the inclusion of the series does not result in a widening of the confidence intervals. In short, the precision with which the break date can be estimated increases with the number of series that have a common break. Also, for a fixed break magnitude, note that the confidence interval does not depend on the sample size; in particular, it does not shrink as the sample size increases. These two facts illustrate the importance of using multiple series in identifying break dates.

[5] A two-sided Brownian motion, $W(\cdot)$, on the real line is defined as $W(v) = W_1(-v)$ for $v<0$, and $W(v) = W_2(v)$ for $v \geqslant 0$, where W_1 and W_2 are two independent Brownian motion processes on $[0, \infty)$, with $W_1(0) = W_2(0) = 0$.

3.2. *Unknown break date, stationary regressors*

The results of BLS apply to general vector autoregressions among stationary variables and also among systems of cointegrated regressors. In this paper, we focus on the stationary case. Specifically, we assume that the errors in the regressions have $4+\kappa$ moments for some $\kappa > 0$. The general form of the regression is (Eq. (2.2) from BLS)

$$y_t = (G_t' \otimes I_n)\theta + d_t(k)(G_t' \otimes I_n)S'S\delta + \varepsilon_t, \tag{6}$$

where y_t is $n \times 1$, G_t' is a row vector containing a constant, lags of y_t, and row t of the matrix of exogenous regressors, X, I_n is an $n \times n$ identity matrix, $d_t(k) = 0$ for $t<k$ and $d_t(k) = 1$ for $t \geqslant k$, and Σ is the covariance matrix of the errors ε_t.[6] θ and δ are parameter vectors with dimension r. For example, for a first-order vector autoregression with a vector of constants μ and parameter matrix A ($y_t = \mu + Ay_{t-1} + \varepsilon_t$), $\theta = vec[(\mu, A)]$ and $r = n(n+1)$. S is a selection matrix containing zeros and ones and having column dimension r and full row rank (equal to the number of coefficients that are allowed to change). It is used to identify (via the placement of the ones) which of the r parameters are allowed to change in the regression. We consider two cases for S. If $S = I_r$, then Eq. (6) is a full structural change model. For $S = s \otimes I_n$, where $s = (1, 0, \ldots, 0)$ is a row vector, Eq. (6) allows for a mean shift only (this is the case considered in the empirical examples of BLS).

Eq. (6) emphasizes the flexibility that any or all of the coefficients may change. We can write the system more compactly as

$$y_t = Z_t'(k)\beta + \varepsilon_t, \tag{7}$$

where $Z_t'(k) = ((G_t' \otimes I_n), d_t(k)(G_t' \otimes I_n)S')$ and $\beta = (\theta', (S\delta)')'$. If we let $R = (0, I)$ be the selection matrix associated with β, then $R\beta = S\delta$ and the F-statistic testing $S\delta = 0$ is

$$\hat{F}_T(k) = T\{R\hat{\beta}(k)\}' \left\{ R \left(T^{-1} \sum_{t=1}^{T} Z_t \hat{\Sigma}_k^{-1} Z_t' \right)^{-1} R' \right\}^{-1} \{R\hat{\beta}(k)\}, \tag{8}$$

where $\hat{\beta}(k)$ and $\hat{\Sigma}_k$ denote the estimators of β and Σ, respectively, evaluated at $\hat{k}$. As before, the statistic of interest is max $F_T(k)$, converging to max F^* (see Eq. (2)). The dimension of F is now equal to the rank of S, which is either r (all coefficients change) or n (only the means change). The corresponding distribution can easily be approximated by partial sums of normal random variables for each dimension; the appendix contains Table 10 with the corresponding critical values for dimensions up to 68. Higher dimensions are available on request.

---

[6] The notation that we employ for $d_t(k)$ here represents a slight departure from the standard break literature; in particular, usually $d_t(k) = \mathbf{1}(t > k)$ where $\mathbf{1}(A) = 1$ when event A is true (and therefore the date k is the last date of the old regime). We choose to adopt the current convention because we believe it is a more intuitive representation of a "break date" (i.e., the date k represents the first date of the new regime).

The dimension of the test statistic increases both with the dimensionality of the system and with the number of regressors in the model whose coefficients are allowed to break. As an example, consider an $n \times 1$ vector autoregression. If the order of the VAR is p and we allow for a break in all of the coefficients, the relevant dimension of the F-statistic will be $n(np+1)$.

To conduct inference about the break date, BLS (Theorem 4) show that analogous to Eq. (3),

$$[\delta_T' S' S(Q \otimes \Sigma^{-1}) S' S \delta_T](\hat{k} - k_0) \overset{\mathrm{d}}{\Rightarrow} V^*,$$

the same limiting distribution as in the case without regressors (with limiting density given by Eq. (4) above) and $Q = \operatorname{plim}(1/T)\sum_{t=1}^{T} G_t G_t'$. Thus we can similarly invert the limiting distribution to construct confidence intervals for the estimated break date, based on allowing any or all of the coefficients to experience a break. The confidence interval is

$$\hat{k} \pm \alpha_{\pi/2}[(S\hat{\delta}_T)' S(\hat{Q} \otimes \hat{\Sigma}_{\hat{k}}^{-1}) S'(S\hat{\delta}_T)]^{-1}, \tag{9}$$

where $\hat{Q} = (1/T)\sum_{t=1}^{T} G_t G_t'$ and $\hat{k}, \hat{\Sigma}_k$ are estimated values. Hence the observations made for the "no-regressor" case still apply when the model is extended to include stationary regressors.

3.3. *Finite sample properties*

Because this is the first paper to implement the multivariate structural break tests allowing for changes in all coefficients, it is important to examine the finite sample properties of the break test statistics and the associated break date confidence interval. BLS investigate the case of a mean break with univariate and trivariate systems and find the tests to have good size and power characteristics. Whether or not these properties hold in more general structural break specifications is the focus of this section.

There are two ways in which the simulations presented here extend the mean-break results of BLS. First, we consider size and power when all the coefficients of the data-generating process and/or the estimation experience a structural change. The test allowing all coefficients to change should have substantially greater power than the test allowing for only a mean break. Second, we introduce exogenous variables (represented by X contained in G_t in Eq. (6)) in the model, allowing for a structural break in the coefficients on these variables (the "world variables", see Section 4).

Table 1 summarizes our Monte Carlo results. We consider univariate systems for equity returns for three countries, bivariate systems for equity returns and dividend yields for Mexico and Thailand, and trivariate systems using equity returns, dividend yields, and the ratio of market capitalization to GDP. The parameters are chosen to reflect actual estimates. We report the empirical size for a 5% break test, the size-adjusted power for that test, and the coverage rate of the confidence interval under the alternative of a break. The empirical size is the percent of replications yielding a test value greater than the 95% critical value of the Sup-Wald test. The null model is

Table 1
Small sample properties of BLS tests

Country	Dependent variable	World instruments included	All coefficients break	Observations	Size (%)	Power (size-adjusted)	Coverage rate (true p)	Empirical critical value (5%)	Lags (p)	Dimension of the test
Mexico	Returns	No	No	244	0.063	0.177	0.764	9.398	1	1
Mexico	Returns	No	Yes	242	0.069	0.074	0.525	16.003	3	4
Mexico	Returns	Yes	No	241	0.065	0.559	0.848	9.458	2	1
Mexico	Returns	Yes	Yes	241	0.105	1.000	0.681	25.489	2	7
Chile	Returns	No	No	244	0.085	0.725	0.837	10.067	1	1
Chile	Returns	No	Yes	243	0.147	0.609	0.702	17.204	2	3
Colombia	Returns	No	Yes	136	0.099	0.077	0.680	13.762	1	2
Colombia	Returns	Yes	Yes	134	0.112	0.985	0.441	22.995	1	6
Mexico	Returns, dividend yields	No	Yes	233	0.181	0.970	0.369	39.931	2	10
Thailand	Returns, dividend yields	No	No	234	0.215	0.098	0.085	16.328	1	2
Thailand	Returns, dividend yields	No	Yes	234	0.231	0.382	0.420	40.093	1	6
Mexico	Returns, dividend yields, market capitalization/GDP	No	Yes	227	0.356	1.000	0.726	80.285	2	21
Korea	Returns, dividend yields, market capitalization/GDP	No	Yes	228	0.412	1.000	1.000	80.526	1	12
Brazil[a]	Returns, dividend yields, market capitalization/GDP	No	Yes	227	0.323	0.823	0.314	80.169	2	21
Jordan[a]	Returns, dividend yields, market capitalization/GDP	No	Yes	201	0.436	0.757	0.553	83.022	2	21
Venezuela[a]	Returns, dividend yields, market capitalization/GDP	No	Yes	119	0.604	0.999	0.754	92.505	2	21

The Monte Carlo analysis uses data generating processes estimated from the data with the order of the VARs(p) determined by the Schwarz (1978) Bayesian Information Criterion. We consider univariate systems for equity returns for three countries, bivariate systems for Mexico and Thailand, and a trivariate system for Brazil, Mexico, Korea, Jordan, and Venezuela. The null model is the autoregressive model estimated over the whole sample with or without instruments added to the regression. The instruments are the lagged world return, the lagged world dividend yield, the lagged Baa-Aaa yield spread, and the lagged change in the slope of the U.S. term structure of interest rates. For the alternative, the model is estimated allowing a break at the break date but constraining the variance–covariance matrix to be the same across the two periods. The empirical size is the percent of replications yielding a test value greater than the 95% critical value of the Sup-Wald test. The power represents the number of rejections under the alternative of a break. The coverage rate determines what percent of the Monte Carlo replications give break estimates that fall into our estimated 90% confidence interval.

[a] 2000 replications used instead of the usual 5000.

simply the vector autoregressive model estimated over the whole sample with or without exogenous instruments added to the regression. These exogenous instruments are the lagged instruments capturing the world business cycle (see Section 5.3). There are four possible tests depending on whether we consider a mean break only in a purely autoregressive system (No–No in Columns 3 and 4), all coefficients changing in a purely autoregressive system (No–Yes in Columns 3 and 4), the mean to break in an autoregressive system with world instruments (Yes–No in Columns 3 and 4), or all coefficients to break in the general system with exogenous instruments (Yes–Yes in Columns 3 and 4).

The power counts the number of rejections at the empirical critical value under the alternative of a break. The data-generating process (DGP) in this case is the VAR estimated accommodating the break tested for under the null. Finally, the coverage rate records the fraction of Monte Carlo replications in which the estimated break date falls in the 90% confidence interval given by Eq. (9). In general, we use 5,000 replications in the Monte Carlo analysis.

Table 1 reports eight different univariate cases. While the tests are well sized when the test allows for a mean break only, there is evidence of some size distortions when the test allows for all coefficients to break and when world instruments are included in the regression. The maximum empirical size is only 14.7%, however. The bivariate systems have an empirical size of about 21.5% and for the trivariate systems, the size distortion is somewhat DGP dependent, varying between 32% and 60%. Not surprisingly, the smaller the sample, the larger is the size distortion. Venezuela has only 119 observations, whereas Jordan has 201 and Mexico, Korea, and Brazil each have 244 observations.

Table 1 demonstrates that the power of the test statistics is high overall, with a few understandable exceptions. In the case of Mexico (univariate), the break magnitude is quite small but the addition of world instruments substantially increases power. Similarly, Colombia is a country with a fairly small sample (131 observations) where the addition of world variables also dramatically increases power. Demonstrating the attraction of adding series that may break at similar dates, the power is exceptionally high for the multivariate systems with power between 75.7% and 100%, except for Thailand. The mean break is simply not large enough in Thailand to yield a powerful test, but allowing all coefficients to break does increase power. Note that the bivariate Mexican system yields very powerful tests without world instruments.

The coverage rates are satisfactory in the univariate cases and very good for Chile and Colombia, when world instruments are included. The latter results are similar to what is reported in BLS for the mean break case only. For the multivariate systems, if we exclude the Thailand case without power, the coverage rates vary between 0.314 and 1.000.

We draw the following conclusions from our Monte Carlo analysis. First, the size properties of the break tests are satisfactory for univariate systems but are often inadequate for multivariate systems, implying overrejection at asymptotic critical values. Therefore, we use empirical critical values to judge significance in the multivariate systems. Second, multivariate systems or systems that include world variables have reasonable to excellent power. Third, the coverage rates of the break intervals are mostly too small, which should be taken into account when interpreting our results.

4. Data

We consider a number of time series for 20 emerging markets followed by the International Finance Corporation (IFC). It is best to think of our variables as constituting five groups: financial data linked to price levels, financial variables related to liquidity, financial flows, financial variables linked to the comovement of returns, and economic indicators.

The first group contains U.S. dollar index total returns and dividend yields from the IFC. Bekaert and Harvey (2000, Section 1) argue that changes in dividend yields in emerging markets are a better proxy for changes in expected returns around liberalizations.

The second group tries to measure the liquidity in the local market. We use value traded to GDP (the annualized dollar volume of trading from IFC divided by nominal GDP in dollars (identical monthly values for each 12-month period from the World Bank). We also use a measure of monthly turnover (12 times monthly dollar volume of trading divided by market capitalization, both from IFC). We also examine the total market capitalization relative to GDP.

The third group focuses on capital flows to the market. We measure the cumulative net U.S. holdings as a percentage of market capitalization in 17 of the 20 emerging markets. Our measure of holdings is based on the net equity capital flows from the *U.S. Treasury Bulletin* (see also Tesar and Werner, 1995). Following Bekaert and Harvey (2000), we allow the value of the holdings to be impacted by both the net flows in a particular period and the realized return on the market (which influences current holdings).

Our fourth group of variables details both the structure and comovement of returns in each market. The stock-to-stock comovement is proxied by the cross-sectional standard deviation of returns each month. If all stocks move together, as would occur in a market with strong sector concentration, the dispersion value is low.[7] Market liberalization could impact the competitive structure of individual markets. To test this, we examine a concentration factor which is a Herfindahl index modified to lie between zero (all firms same size) and one (one dominant firm):

$$CF = \sqrt{\left(\frac{N}{N-1}\right)\sum_{i=1}^{N}\left(w_i - \frac{1}{N}\right)^2},$$

[7] Christie and Huang (1995) use the cross-sectional standard deviations of U.S. securities to measure herding. Bessimbinder et al. (1996) examine dispersion in the context of trading volume. Bekaert and Harvey (1997) use the dispersion measure to help explain the cross-section of volatility in emerging markets. See also Connolly and Stivers (2001).

where $N > 1$ represents the number of firms at a point in time for a particular country and w_i is the share of capitalization for firm i.

The fourth group also incorporates measures of comovement with world variables. We report the conditional betas and correlations with world market returns before and after breaks. These conditional moments are based on the multivariate GARCH model presented in Bekaert and Harvey (1997). This model allows for time-variation in expected returns, asymmetry in volatility, and non-normalities. Most important, the emerging market expected return is determined by a combination of world information and local information, with the importance of the world information increasing with the degree of integration. Similarly, the local volatility is a function of both local shocks and world shocks, with the world shocks being more important in integrated markets. This model produces monthly fitted correlations and betas for our analysis. We also report a measure of ex post volatility based on three-year rolling standard deviations estimated from the returns data.

Our final group of variables provides information on the local economic environment. Both the inflation rates and foreign exchange volatility (standard deviation of monthly changes in the foreign exchange rate to the dollar over the past three years) represent the stability of the monetary policy. The degree of economic integration is proxied by the size of the trade sector (exports plus imports divided by GDP). As an ex ante measure of a country's prospects and of the credibility of policies, we also examine the Institutional Investor country credit rating which is available on a semiannual basis. Erb et al. (1996) show that the cross-section of credit ratings is correlated with expected returns and volatility in world markets. We also include the World Bank's measure of the real exchange rate (index of trade-weighted exchange rates adjusted for inflation). Finally, we consider real per capital GDP growth as reported by the World Bank.

5. Empirical results

With some 11 time series and 20 countries, there are a huge number of systems that we can estimate. Section 5.1 summarizes the results for univariate break tests on all series in all countries allowing for a mean break only, whereas Section 5.2 focuses on univariate models allowing for breaks in all parameters. In Section 5.3, we assess the impact of adding exogenous world variables. For the multivariate systems in Section 5.4, we focus on financial variables that play a role in most models of financial market integration: returns, dividend yields, capital flows, and the ratio of market capitalization to GDP. Section 5.5 repeats part of our analysis for developed markets as a control experiment. Section 5.6 groups countries in one system for one variable rather than multiple variables for one country.

5.1. Univariate analysis: mean breaks

Our univariate analysis begins with a country-by-country analysis of 11 time series. These include: U.S. dollar returns, dividend yields, market capitalization to

GDP, turnover, value traded to GDP, net equity capital flows, concentration ratios, cross-sectional standard deviations, change in inflation rates, foreign exchange volatility, and Institutional Investor's country credit ratings.

The estimation involves univariate autoregressions with the lag length determined by the Bayesian Information Criterion of Schwarz (1978). Initially, this analysis only allows for breaks in the mean of the series. That is, the autoregressive parameters are assumed to be constant across the break point. While this assumption is not particularly attractive, it gives us a base case to work with.

Our results can be briefly summarized with three observations. First, with few exceptions, the break point for U.S. dollar returns has a wide confidence interval, often wider than the sample. Second, for all other series, between 50% and 75% of the countries show significant breaks at the 5% confidence level. This even occurs for the macro-level data and credit ratings which are not available at a monthly frequency.[8] Third, there might be significant breaks for only a few series for any particular country, reinforcing the need to examine multiple series simultaneously. For example, Venezuela, a country that did not show significant breaks with other economic series, presents a highly significant break in net U.S. holdings in March 1994. In multivariate analysis below, we detect significant break points in financial series as well.

5.2. *Univariate analysis: breaks in all parameters of autoregression*

While others have examined structural breaks in means or trends (e.g., Zivot and Andrews, 1992; Ben-David and Papell, 1998), this is the first paper to empirically test for breaks in all of the parameters of the autoregression.[9] Examining breaks in 11 series in over 20 countries again generates too much information to be fully reported in tables and graphs. In this section, we illustrate the main points. Full results are available on the Internet at http://www.duke.edu/~charvey/Research/inder.htm.

First, an examination of the test statistics indicates that it is much more likely that the breaks are statistically significant when we consider breaks in all the parameters rather than the mean alone. To highlight this point, Table 2 presents the analysis of one critical financial series, U.S. dollar equity returns in 20 different markets. As we have indicated before, we expect financial market integration to lower expected returns. While it is extremely difficult to determine breaks in the mean of equity returns (in the first panel of Table 2), there are more countries with significant breaks when all the parameters are allowed to change (second panel of Table 2). For example, the second panel suggests a significant break for Chile and Mexico whereas the evidence in the first panel provides weak or no support for the existence of these breaks.

Second, since the noisiness of returns prevents us from finding significant breaks for the majority of the countries, we must supplement the information in returns

[8] The credit rating series covers a short time span and has a semiannual frequency. Therefore, the results for this series cover only 22 observations. Results may be sensitive to this very small sample.

[9] Note that we still assume a constant innovation covariance matrix across the break.

Table 2
Analysis of structural breaks in emerging market equity returns

	A. Mean break			B. All parameters break			C. All parameters break + 4 instruments		
Country	5th percentile	Median	95th percentile	5th percentile	Median	95th percentile	5th percentile	Median	95th percentile
Argentina		Jun-85		Feb-84	Jul-89	Dec-94	May-79	Jun-79	Jul-79
Brazil		Sep-83	Aug-94		Sep-83	Aug-94	Dec-86	Jan-88	Feb-89
Chile	Mar-78	Jul-80**	Nov-82	Jan-79	Nov-79***	Sep-80	Apr-80	May-80***	Jun-80
Colombia	Jun-90	Apr-94		Aug-90	Feb-92**	Aug-93	Jan-92	Feb-92**	Mar-92
Greece	Aug-80	Nov-85*	Feb-91	Nov-88	Aug-90*	May-92	Apr-85	May-86**	Jun-87
India	Mar-85	Apr-92		Oct-84	Jun-90	Feb-96	Jul-90	Aug-90**	Sep-90
Indonesia	Jan-91	Nov-91**	Sep-92	Jan-91	Nov-91**	Sep-92	Sep-91	Nov-91	Jan-92
Jordan	Oct-79	Feb-82*	Jun-84	Oct-79	Feb-82*	Jun-84	May-80	Mar-82	Jan-84
Korea	Mar-81	Apr-89		Mar-81	Apr-89		Dec-80	Jan-81	Feb-81
Malaysia		Dec-87		Jul-91	Jan-94		Dec-86	Jan-87**	Feb-87
Mexico		Jan-83	Oct-94	Oct-85	Oct-87***	Oct-89	Mar-86	May-86*	Jul-86
Nigeria	Apr-89	Apr-93		Apr-89	Apr-93		Dec-93	Jan-94***	Feb-94
Pakistan	Jul-91	Mar-94		Jan-92	Dec-93	Dec-95	Dec-91	Jan-92*	Feb-92
Philippines	Oct-86	Aug-87***	Jun-88	Oct-86	Aug-87***	Jun-88	Dec-92	Jan-93	Feb-93
Portugal	Jun-87	Feb-88***	Oct-88	Nov-87	Jan-88***	Mar-88	Nov-87	Dec-87***	Jan-88
Taiwan	Jul-86	Jun-89	May-92	Jul-86	Jun-89	May-92	Jun-92	Feb-93*	Oct-93
Thailand		Jun-86		Mar-85	Oct-87	May-90	Mar-79	May-79	Jul-79
Turkey		Aug-90	Oct-94		Aug-90	Oct-94	Aug-89	Oct-89	Dec-89
Venezuela	Dec-85	Feb-92		Jun-90	Feb-92**	Oct-93	Apr-89	Feb-90*	Dec-90
Zimbabwe		Jul-84		Aug-80	Oct-84	Dec-88	Jul-91	Aug-91**	Sep-91

The estimation involves univariate autoregressions with the lag length determined by the Schwarz (1978) Bayesian Information Criterion. In Panel A, only the mean is allowed to change (i.e., the parameters in the autoregression are constant across the break point). In Panel B, all of the parameters of the autoregression are allowed to change. In Panel C, we introduce four instrumental variables: the lagged world return, the lagged world dividend yield, the lagged Baa-Aaa yield spread, and the lagged change in the slope of the U.S. term structure of interest rates. We report the median break point as well as the 90% confidence interval for the break. *, **, *** indicate significance levels of 10%, 5% and 1%, respectively, for the Sup-Wald statistic which measures whether the break is significant. If the date is left open, it is outside the sample interval.

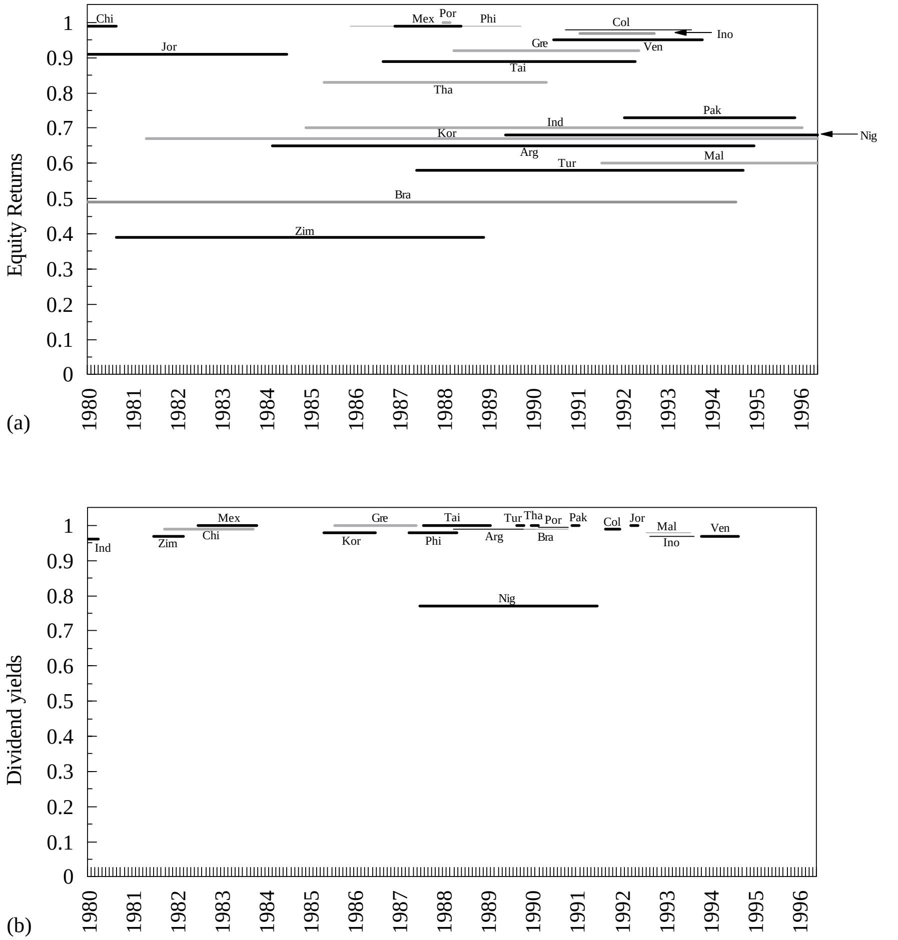

Fig. 1. Structural breaks by variable. The 90% confidence interval around the break date (horizontal axis) from the univariate break analysis is graphed against the confidence that a break occurred (one minus the p-value from the statistical break test; vertical axis) for two variables: returns and dividend yields, presenting all countries in each panel.

with other variables. Bekaert and Harvey (2000) present simulation evidence which suggests that it is easier to detect a break in the cost of capital by examining dividend yields rather than returns. Fig. 1 provides considerable supporting evidence for their simulations. In the two panels of Fig. 1, we graph the 90% confidence interval around the break date (horizontal axis) against the confidence that a break occurred (one minus the p-value from the statistical break test; vertical axis) for returns and dividend yields, putting all countries on one graph.[10] Notice that the line for the dividend yield is higher (significant break) and shorter (tight confidence interval)

[10] A full set of figures for all series (Fig. 1) and all countries (Fig. 2) is available on the Internet.

than the line for returns in all countries. The breaks for dividend yields are much more likely to be significant and the confidence intervals are tighter. We combine the information in the dividend yields and returns in a multivariate analysis in Section 5.4.

Third, the break dates for other variables related to market integration and stock market development, such as market capitalization to GDP, turnover, value traded to GDP, and cumulative net capital flows to market capitalization, are much more clustered in time across countries than the break dates for the returns series (all figures are available on the Internet). They also occur generally later than for most other series. One potential explanation for this phenomenon is that the capital flows (and any market integration they bring about) are primarily driven by factors in the developed world, such as interest rates (see World Bank, 1997; Stulz, 1999). Interest rates were unusually low in the U.S. in 1993, which coincided with a large capital outflow to emerging markets. In Section 5.4, we add either market capitalization to GDP or both market capitalization to GDP and the capital flows series to the returns–dividend yield system to investigate this further. In Section 5.5, we also consider grouping across countries rather than across variables.

Fourth, Fig. 2 examines the break dates for the 11 series for three countries illustrating three representative patterns that we find. The depiction is analogous to Fig. 1 but groups the breaks (plus interval) for all series in one figure for three countries: Chile, Colombia, and Korea.

Let us first focus on Colombia, for which the break dates for the various series appear clustered in time. This is consistent with a far-reaching event such as market integration driving the breaks in all series. An analysis of the economic reforms suggests that integration occurred in the early 1990s. The analysis of the break in means provides inconclusive results for both returns and dividend yields (no significant break). In the figure (where all parameters are allowed to break), we see a significant break in U.S. dollar returns in February 1992. The analysis of dividend yields suggests a slightly earlier date, October 1991. Around this time we also observe a significant shift in market capitalization to GDP, turnover, and credit ratings. The grouping of this information should further help pin down the date of integration.

For the majority of countries, this clustering does not happen and there seem to be two focal points for the break dates. One of the series that often breaks early on is the credit rating series. It is likely that for countries such as Argentina, Brazil, Chile, Mexico, and Venezuela, the first clustering represents the start of the debt crisis in 1982. The second clustering potentially reflects the onset of the market integration process. As Fig. 2 illustrates for Chile, virtually all variables break in the early 1980s, with the exception of the stock market development variables and the capital flow series which break in the 1990s.

Korea represents a third pattern. Fig. 2 shows that different series break at different times and many of the break dates have large confidence intervals. This is not surprising since Korea experiences a gradual liberalization process. For such a country, it will be particularly important to pool information from different series.

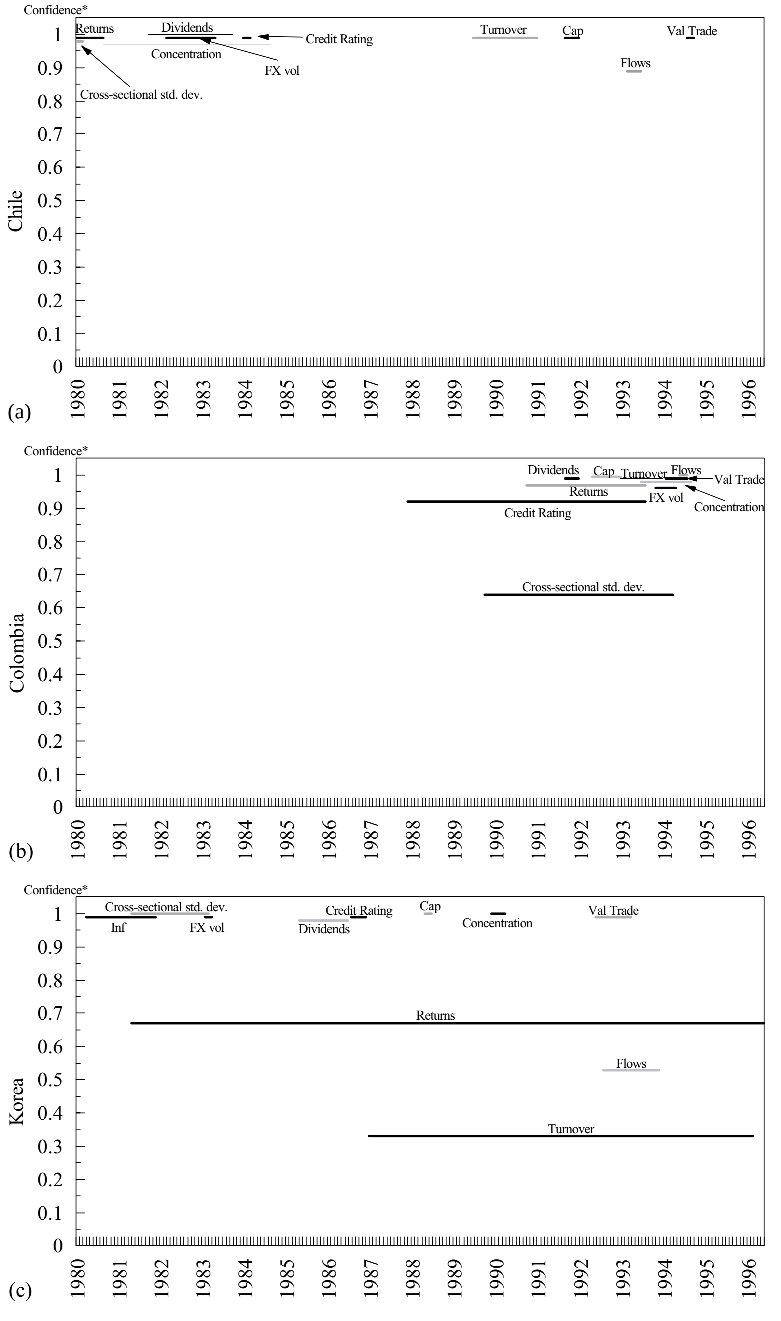

Fig. 2. Structural breaks by country. The 90% confidence interval around the break date (horizontal axis) from the univariate break analysis is graphed against the confidence that a break occurred (one minus the *p*-value from the statistical break test; vertical axis) for three countries: Chile, Colombia and Korea, presenting all variables in each panel. The variables include returns, dividend yields, market capitalization to GDP, turnover, value traded to GDP, U.S. holdings, concentration ratios, cross-sectional standard deviation, inflation, and trailing three-year FX volatility.

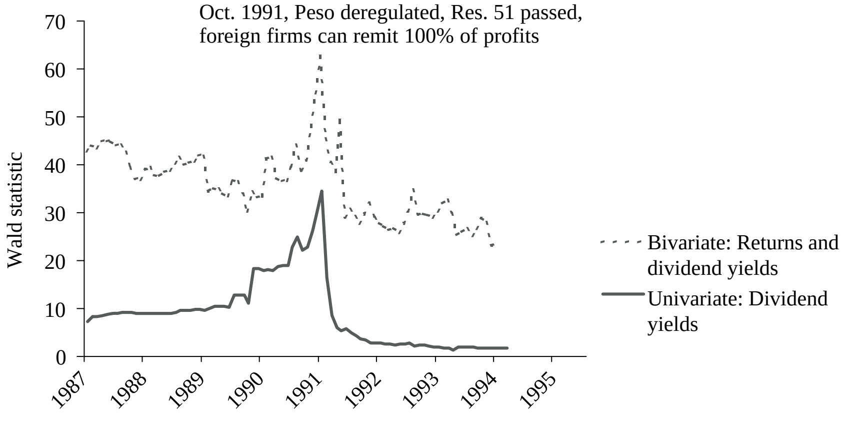

Fig. 3. Time series of Wald statistics for Colombia. We report Wald statistics from a univariate model for dividend yields and a bivariate model for returns and dividend yields where all coefficients are allowed to break.

It is a limitation of our analysis that we identify only one break in each time series. By calculating the Wald statistic at every point in the time series, we can informally assess the appropriateness of this assumption. For example, Fig. 3 presents the Wald statistics for the univariate analysis of dividend yields for Colombia. The Wald statistic moves above the critical value in late 1990, which coincides with a significant number of reforms taking place in that market. There is no evidence for Colombia of a second break.[11]

5.3. *The importance of world factors*

The third panel of Table 2 brings world information into the univariate autoregression analysis. As in the other panels of Table 2, all parameters of the model are allowed to change permanently at the break point. The intuition for inserting the world variables and allowing for their coefficients to change is that world factors are likely to be more influential after capital market liberalizations (see Bekaert and Harvey, 1997). For example, discount rates should reflect global, rather than local, systematic risk after liberalization.

We augment the autoregressions with the lagged world return and three additional instruments that contain information about the world business cycle: the lagged world dividend yield, the lagged Baa-Aaa yield spread, and the lagged change in the slope of the U.S. term structure of interest rates. The confidence intervals shrink

[11] Some caution needs to be exercised in interpreting the Wald statistic plots. In particular, the values of the statistic are serially correlated, so that, for example, if the p-value corresponding to a test for a break at point x is half as large in magnitude as the analogous test for a break at point y, it does not necessarily mean that it is twice as likely that a break occurred at point x relative to point y. In addition, we emphasize that our graphical discussion does not constitute a consistent statistical test for multiple breaks (instead, see Bai and Perron, 1998), although we believe it provides a reasonable characterization of the likelihood of breaks in the series at different dates.

quite dramatically. For example, the 90% confidence interval for the break date in Colombia (February 1992) is much tighter than in the second panel, spanning one month before and one month afterwards. Whereas for most countries the wide interval resulting from the returns analysis is simply shrunk by adding the world information, for some other countries the new date is outside the wide confidence interval (see, for instance, Argentina and Zimbabwe).[12]

5.4. *Multivariate analysis*

Table 3 compares the results of univariate and multivariate estimates to some predetermined dates presented in Bekaert and Harvey (2000). The univariate analysis examines breaks in the ratio of U.S. holdings to market capitalization, detecting when U.S. capital flows first substantially increase. The multivariate results begin with two bivariate systems: returns plus dividend yields and returns plus market capitalization to GDP. Although both dividend yields and market capitalization to GDP can capture permanent price effects, they might not break at the same time (see also Fig. 1). As with the results in the final panel of Table 2, we include four world instruments in this analysis. In addition, we allow for all of the parameters in the VAR to break. This includes the parameters on the world instruments.

The bivariate analysis allows us to more accurately detect break points in the financial series compared to the univariate estimation. In the analysis of returns and dividends, 18 of the 20 countries have breaks that are significant at the 1% level.[13] In the analysis of returns and market capitalization to GDP, we detect significant breaks at the 1% level in 19 of the 20 countries and one country has a break that is significant at the 5% level. Generally, the confidence intervals around the break dates are only two months wide, although this interval may somewhat underestimate the 90% interval (see Section 3.3). Interestingly, in 16 of the 20 countries, the break date in the system with market capitalization to GDP is later than in the system with dividend yields. We will return to this below.

There is the possibility that the breaks are only asymptotically significant given the overrejection we observe for the bivariate systems in our Monte Carlo analysis. However, for the market capitalization systems, the smallest statistic is 59.44 for $p = 2$, much larger than the bivariate empirical critical value of Table 1. For the dividend yield system, a few countries record statistics between 32 and 40, which are likely only marginally significant.

The bivariate return and dividend yield system is added to the univariate analysis of dividend yields presented for Colombia in Fig. 3. For Colombia, there is one clear break that is obvious from the time series of Wald statistics, in October 1991. This

[12] Given that the confidence intervals constitute 90% of the distribution, this is not unexpected in a sample of 20 countries.

[13] We ran into estimation problems for Indonesia due to its small sample and we report results only for the case that does not allow the world instruments to break.

Table 3
Predetermined and estimated break point analysis with world instruments

Country	Official liberalization date	ADR introduction	Country fund introduction	Univariate: (1) holdings/mkt.cap. Estim. break	Bivariate: (1) returns, (2) dividend yield Estim. break	Bivariate: (1) returns, (2) mkt.cap./GDP Estim. break	Trivariate: (1) returns, (2) dividend yield, (3) mkt.cap./GDP Estim. break	Quadravariate: (1) Returns, (2) dividend yield, (3) mkt.cap./ GDP, (4) holdings/ mkt.cap. Estim. break
Argentina	Nov-89	Aug-91	Oct-91	May-93***	Feb-89***	Aug-91***	Aug-91***	Jun-92***
Brazil	May-91	Jan-92	Oct-87	Jun-92***	Apr-90***	Jan-93***	Apr-90***	Apr-90***
Chile	Jan-90	Mar-90	Sep-89	Apr-93	Feb-83***	Jan-93***	Feb-83***	Jan-93***
Colombia	Feb-91	Dec-92	May-92	May-94***	Oct-91***	Jan-93***	Oct-91***	May-94***
Greece	Dec-87	Aug-88	Sep-88	Dec-86***	Jun-86***	Apr-90***	Dec-90***	Aug-90***
India	Nov-92	Feb-92	Jun-86	May-93***	Nov-87***	Jan-92***	Jan-92***	May-93***
Indonesia[a]	Sep-89	Apr-91	Jan-89	Feb-95	Aug-92***	Feb-95**	Nov-93***	Aug-93***
Jordan	Dec-95	n/a	n/a	No holdings data	Apr-92***	Apr-88***	Apr-92***	No holdings data
Korea	Jan-92	Nov-90	Aug-84	Mar-93*	Feb-80***	May-88***	May-88***	Sep-88***
Malaysia	Dec-88	Aug-92	Dec-87	Feb-93	Jan-93***	Oct-93***	Jul-88***	Jan-94***
Mexico	May-89	Jan-89	Jun-81	Jan-92**	Feb-83***	Jan-93***	Feb-83***	Jan-92***
Nigeria	Aug-95	n/a	n/a	No holdings data	Jun-89	Jan-94***	Jan-93***	No holdings data
Pakistan	Feb-91	n/a	Jul-91	Oct-94***	Dec-90***	Oct-93***	Dec-91***	Oct-94***
Philippines	Jun-91	Mar-91	May-87	Jan-90***	Oct-87***	Oct-93***	Dec-93***	Jan-90***
Portugal	Jul-86	Jun-90	Aug-87	Sep-94	Aug-88***	Dec-87***	Jul-88***	Jun-88***
Taiwan	Jan-91	Dec-91	May-86	May-94	Aug-88***	Jan-88***	Oct-88***	Oct-88***
Thailand	Sep-87	Jan-91	Jul-85	Jul-88***	Jan-90**	Jan-93***	Mar-93***	Oct-93***
Turkey	Aug-89	Jul-90	Dec-89	Mar-94	Jun-89***	Feb-91***	Sep-89***	May-89***
Venezuela	Jan-90	Aug-91	n/a	Feb-94***	Jan-94***	Jan-92***	May-94***	Jan-94***
Zimbabwe	Jun-93	n/a	n/a	No holdings data	Mar-83	Nov-91***	Aug-90***	No holdings data

The first three columns are from Bekaert and Harvey (2000). The next five columns result from (vector) autoregressions with the lag length determined by the Schwarz (1978) Bayesian Information Criterion. We consider up to four lags for the VAR, with the exception of the quadravariate systems where we consider at most three lags. The first panel is a univariate result. The next two panels present the bivariate results. The last columns present the trivariate and quadravariate results. The bivariate systems also include world variables but the trivariate and quadravariate systems do not. In all multivariate estimations, the 90% confidence interval for the break was one month before and one month after the median break. *, **, *** indicate significance levels of 10%, 5% and 1%, respectively, for the Sup-Wald statistic which measures whether the break is significant.

[a] Estimation for Indonesia does not allow for a break in the world instruments.

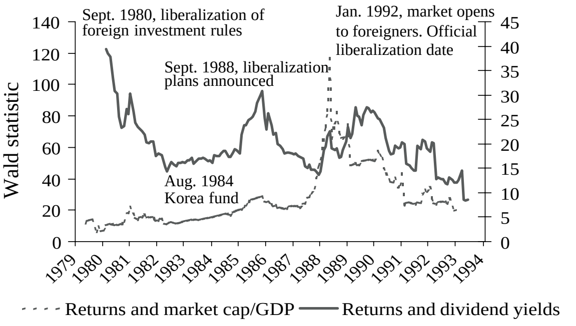

Fig. 4. Time series of Wald statistics for Korea. We report Wald statistics from bivariate models for returns and dividend yields (right axis) as well as returns and market capitalization to GDP (left axis) where all coefficients are allowed to break.

month exactly coincides with major regulatory initiatives that impacted foreign exchange conversion and the ability of corporations to remit profits abroad. On the other hand, we know that Korea experienced multiple liberalizations. Fig. 4 presents the Wald statistics for both the bivariate returns and dividend yield as well as the returns and market capitalization to GDP system. There are roughly four peaks in the graph of the Wald statistic for returns and dividend yields. There is some correspondence between these peaks and economic events. For example, in September 1980 there was a liberalization of rules for foreign investment in Korea (first peak). This opened the market to foreign direct investment. In August 1984, the Korea Fund was launched. This gave foreign investors their first chance to make portfolio investments in Korea (second peak). In fact, when we drop the default and term spread instruments, the break date becomes November 1985, which is significant at the 5% level. It is possible that the early break in 1980 is caused by breaks in these two U.S. instruments (see below). In September 1988 (third peak), sweeping liberalization plans were announced (that were not implemented until much later).

When dividend yields are replaced by market capitalization to GDP in the bivariate system, the resulting break date is May 1988 (see Table 3 and Fig. 4). The statistics are above their critical values between October 1987 and January 1991, with a spike in May 1988. The only other time that the statistic is above its critical value is January 1992, which coincides with the "official" liberalization date. Indeed, in September 1991, the government announced that the stock market would become "open" to international investors in January 1992 (fourth peak). This meant that foreigners could own up to 10% of the capitalization of a company and no individual could own more than 3%. However, a number of companies had more than 10% foreign ownership. The government had to raise the total foreign ownership limit to 25% for 45 firms. In December 1994, the government raised the foreign ownership limit to 12% from 10%. They also announced their intentions of raising the limit to 15% some time in 1995.

The announcement date for the 15% limit was July 1995.[14] In summary, there are strong a priori grounds for believing that Korea has experienced multiple breaks.

Given that the market capitalization to GDP and dividend yield systems often produce different dates, we consider a trivariate system with all three variables—returns, market capitalization to GDP, and dividend yields—but without the world variables. The seventh column in Table 3 reports the results. Now, every country has a break that is statistically significant at the 1% level and a break confidence interval of two months. When taking small-sample overrejection into account, the evidence for breaks in some countries is weaker. For example, Korea and Venezuela have Sup-Wald statistics that are less than the empirical critical values (5% level) detailed in Table 1. While our Monte Carlo analysis did not consider Indonesia and Malaysia, there is good reason to believe that the breaks in these countries are not strongly significant. In nine of the 20 countries, the dividend yield date is the break date chosen by the trivariate system, and in seven cases it is the market capitalization to GDP date or close to this date. In three cases, the date is between the dates in the bivariate systems, perhaps suggesting a long period of reforms with differently timed effects on the variables we examine.

The most surprising result is that the trivariate system selects an entirely new date for Malaysia, namely July 1988, whereas the dates we find in bivariate analysis are both in 1993. However, further analysis reveals that there was, even then, a local peak in the Wald statistic around the new date.[15]

In the final column of Table 3, we report the break dates for the quadravariate system, adding the capital flows series (cumulative U.S. holdings to GDP) to the trivariate system. Since we expect equity market integration to take effect as soon as the marginal investor changes to being a foreign investor, this may be a powerful variable to tie down the true integration date. Of course, market returns (and hence dividend yields and the market capitalization to GDP variables) might anticipate such capital inflows and break sooner. In the 17 countries with holdings data, we find nine break dates that overlap with the confidence intervals in the bivariate and trivariate analysis. In six countries, the quadravariate date is very close to the confidence intervals estimated previously (within one year). The two remaining countries (Argentina and Colombia) have a break date more than 12 months different from the bivariate and trivariate analysis. This date is also later than the break date identified by other systems.

---

[14] The ownership limit was subsequently raised to 18% in May 1996. In September 1996, the limit was raised to 20%. In May 1997, the ownership limit was further increased to 23%. It was abolished after the Asian crisis.

[15] December 1988 is the official liberalization date reported by the IFC. The liberalization activity clustered in 1992–94 includes August 1992, the announcement of the first Malaysian ADR, and 1993, with the 30% ownership limit on manufacturing firms lifted and a computerized trading system introduced.

5.5. A control experiment using world returns and world dividends

The bivariate analysis in Table 3 assumes that there are breaks in the relations between the country variables and the world information. One might conjecture that some of the breaks we detect are a result of spurious breaks in the process generating the world instruments themselves. In addition, structural breaks can occur for reasons that have nothing to do with market integration. For example, our specification captures time-variation in expected returns through a stable relation between returns and instruments such as dividend yields. Breaks in this relation might have occurred across the world.[16]

As a control experiment, we test whether there are any significant breaks in the world variables. These results are presented in Table 4. We fail to detect a significant break in the univariate return specifications, but we do detect a significant break in the dividend yield regression when world instruments are used as regressors. The same break materializes in the analogous bivariate specification that also includes world returns as an independent variable. Further analysis reveals that the break is in fact caused by a break in the corporate bond spread and term to maturity spread instruments. The break date of mid-1980 follows the change in the operating procedure of the Federal Reserve in the U.S.[17] Given that this break occurs before all of the breaks we find in emerging markets (with one exception, Korea, see above) and even before many of our samples start, it will not affect our analysis.[18]

Overall, this analysis gives us some confidence that when we add the world variables to the analysis, we should be detecting breaks that result from changing relationships to the world variables rather than shifts in the world aggregates alone. Moreover, the significant breaks we find in emerging markets suggest structural changes that apparently did not occur in the developed world.

5.6. Multicountry analysis

The possibility that exogenous events such as low world interest rates, or a relaxation of foreign investment restrictions on the portfolios of institutional investors, will drive portfolio flows into emerging markets implies that it also makes sense to group countries rather than variables to find a common break date.[19] We perform such an analysis in Table 5. Unfortunately, given the sample sizes available to us, we can only investigate four countries and one variable at a time. We examine Latin America (Argentina, Brazil, Chile, and Mexico) and Southeast Asia (Korea, Malaysia, Taiwan, and Thailand).

---

[16] See Garcia and Ghysels (1998) and Bossaerts and Hillion (1999).

[17] The Federal Reserve Bank adopted a "non-borrowed reserve" policy in October 1979 which contributed to significant interest rate volatility. This new policy was abandoned in October 1982.

[18] Of course, many other countries in the world market (such as Japan and New Zealand) experienced major changes in capital controls in the early 1980s. However, at the time, the world market portfolio was dominated by the capitalization of the U.S. equity market.

[19] We thank the referee for suggesting this analysis.

Table 4
A test of whether there are structural breaks in the world variables

Variable	World instruments	Description	Break analysis		
			5th percentile	Estimate	95th percentile
Univariate: World return	No	Coefficients on lags of world return allowed to break	Sep-87	Oct-90	Nov-93
Univariate: World return	Yes	Coefficients on lags of world return and the three instrumental variables allowed to break	Jul-84	Oct-85	Jan-87
Univariate: World dividend yield	No	Coefficients on lags of world dividend yield allowed to break	Nov-74	Apr-80	Sep-85
Univariate: World dividend yield	Yes	Coefficients on lags of world dividend yield and the three instrumental variables allowed to break	May-80	Jun-80***	Jul-80
Bivariate: World returns, dividend yield	No	Coefficients on lags of world return and dividend yield allowed to break	Jul-82	Aug-82	Sep-82
Bivariate: World returns, dividend yield	Yes	Coefficients on lags of world return, world dividend yield and the two instrumental variables allowed to break	May-80	Jun-80***	Jul-80

The estimation involves univariate (bivariate) autoregressions (vector autoregressions) with the lag length determined by the Schwarz (1978) Bayesian Information Criterion. In all tests, all of the parameters of the autoregression are allowed to change. We examine cases with and without four instrumental variables: the lagged world return, the lagged world dividend yield, the lagged Baa-Aaa yield spread, and the lagged change in the slope of the U.S. term structure of interest rates. We report the estimated break point as well as the 90% confidence interval for the break. *, **, *** indicate significance levels of 10%, 5% and 1%, respectively, for the Sup-Wald statistic which measures whether the break is significant.

Table 5
Multicountry break point analysis

Country	Returns Estimated break	Dividend yields Estimated break	Mkt.cap./GDP Estimated break	Holdings/mkt.cap. Estimated break
Latin America Mexico, Argentina, Brazil, Chile	May-82***	Feb-90***	Jan-93***	Jun-92***
Asia: Thailand, Malaysia, Korea, Taiwan	Mar-88***	Feb-91***	Oct-93***	Sep-88***

The estimation involves vector autoregressions with the lag length determined by the Schwarz (1978) Bayesian Information Criterion. The maximum lag we consider is $p = 3$, which is also the order chosen for all the VARs. In all estimations, the 90% confidence interval for the break was one month before and one month after the estimated break. *, **, *** indicate significance levels of 10%, 5% and 1%, respectively, for the Sup-Wald statistic which measures whether the break is significant.

There are a number of interesting results. First, the grouping procedure yields highly significant breaks, even for the return series. For Latin America, we find dates that are quite far apart across the different variables, whereas for Southeast Asia, the dates are more closely clustered. This is perhaps the opposite of what we anticipated, since the recent literature on the Mexican currency crisis has stressed the importance of U.S. capital flows in Latin America. It appears that the effect of the onset of the debt crisis on equity returns was dominant, yielding a 1980 break date. Whereas for two of the four countries (Chile and Mexico) dividend yield systems produce break dates around the start of the debt crisis, the inclusion of Argentina and Brazil changes the break date to early 1990, close to the capital market reforms in all four countries. The market capitalization to GDP and the holdings break dates are later, January 1993 and June 1992, respectively. As is well known, at the end of 1992 and especially in 1993, U.S. investors poured billions of dollars into emerging markets and this exogenous effect could be what is driving this finding.

In Southeast Asia, returns break in March 1988, dividend yields in February 1991, and U.S. holdings to GDP in September 1988. The break date for market capitalization to GDP occurs last in October 1993. The early dates are not so surprising, since both Korea and Taiwan have 1988 break dates in Table 4, and Malaysia shows a break date in 1988 in the quadravariate system in Table 4. Moreover, capital market reforms in Thailand happened in 1987–88 and the capital flow break (see Fig. 1) occurs in July 1988. The 1991 date for dividend yields is in between the break dates for the bivariate return–dividend yield systems for Malaysia and Thailand. The latter date for the bivariate market capitalization variable reflects the fact that we also find a late date for the market capitalization system for Thailand and Malaysia in Table 3.

6. Interpretation

6.1. Did liberalization occur and what does it mean?

Table 6 presents an analysis of the behavior of financial and economic aggregates before and after a break. For this analysis, we estimate the following regression:

$$Y_{t,i} = a_i + \gamma\, Trans_{t,i} + \theta\, Lib_{t,i} + \varepsilon_{t,i}, \tag{10}$$

where a_i is a fixed effect, $Trans_{t,i}$ is a dummy variable equal to one around the liberalization date (two months before the liberalization month and two months after) and $Lib_{t,i}$ is set equal to one after liberalization (starting three months after). Given that many liberalizations occur towards the end of the sample, we are somewhat limited in the number of horizons we can consider. We consider a horizon of five years before and after (a total of 125 observations per country, including the transition period) and a horizon of three years before and after (a total of 77 observations). The average number of observations available per country depends on the variable used and the break date considered and varies between 81.3 (56.5) at a

Table 6
Response of economic variables to liberalization measures (60-month before/after windows, five-month transition window)

	ADR introduction	Official liberalization	Country fund intro	Univariate: holdings/ mkt.cap.	Bivariate 1: returns and dividend yields	Bivariate 2: returns and mkt.cap./GDP	Trivariate	Quadravariate
Market Integration								
Returns	−0.89	0.15	0.14	−0.62	0.26	−1.46	−0.83	−2.07
Std. error	*0.47*	*0.43*	*0.45*	*0.48*	*0.43*	*0.43*	*0.43*	*0.50*
Dividend yields	−1.36	−2.55	−2.98	−1.23	−2.09	−1.26	−1.64	−1.23
Std. error	*0.09*	*0.10*	*0.11*	*0.09*	*0.11*	*0.09*	*0.10*	*0.09*
Mkt.cap./GDP	10.09	10.06	6.72	8.77	5.61	15.13	11.07	11.56
Std. error	*0.54*	*0.47*	*0.31*	*0.60*	*0.32*	*0.54*	*0.41*	*0.57*
Holdings/mkt.cap.	2.07	1.91	2.58	3.46	1.32	0.58	0.84	2.31
Std. error	*0.11*	*0.10*	*0.09*	*0.13*	*0.09*	*0.12*	*0.10*	*0.13*
Beta	0.22	0.22	0.08	0.25	0.07	0.22	0.14	0.18
Std. error	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*
Correlation	0.08	0.08	0.04	0.08	0.02	0.09	0.05	0.07
Std. error	*0.005*	*0.005*	*0.004*	*0.006*	*0.004*	*0.003*	*0.004*	*0.005*
Ex post volatility	−0.58	2.43	3.27	1.13	2.98	−0.23	−1.21	0.04
Std. error	*0.74*	*0.69*	*0.67*	*0.78*	*0.68*	*0.72*	*0.70*	*0.83*
Stock market development								
Turnover	7.53	9.36	24.04	6.43	8.82	9.46	5.17	6.93
Std. error	*1.30*	*1.24*	*1.20*	*1.39*	*1.22*	*1.16*	*1.16*	*1.28*
Value traded/GDP	3.58	2.54	3.27	4.86	2.71	6.28	4.33	5.94
Std. error	*0.35*	*0.29*	*0.19*	*0.36*	*0.32*	*0.29*	*0.29*	*0.36*

Table 6 (*continued*)

	ADR introduction	Official liberalization	Country fund intro	Univariate: holdings/ mkt.cap.	Bivariate 1: returns and dividend yields	Bivariate 2: returns and mkt.cap./GDP	Trivariate	Quadravariate
Concentration ratio	−1.58	−0.68	−4.36	−1.18	−1.71	−0.46	−1.90	−1.48
Std. error	*0.22*	*0.24*	*0.25*	*0.27*	*0.25*	*0.25*	*0.23*	*0.23*
Cross-section std. dev.	−1.04	0.59	1.19	−0.20	1.16	−0.58	−0.40	−1.33
Std. error	*0.28*	*0.28*	*0.29*	*0.30*	*0.29*	*0.27*	*0.28*	*0.30*
Macroeconomic variables								
Credit rating	5.77	3.86	4.73	4.76	0.94	3.93	0.98	7.04
Std. error	*0.93*	*0.86*	*0.88*	*1.08*	*1.03*	*0.75*	*0.95*	*0.93*
FX volatlity	−0.79	−0.41	−0.98	−0.62	0.86	0.82	0.92	−0.41
Std. error	*0.11*	*0.12*	*0.11*	*0.12*	*0.12*	*0.11*	*0.12*	*0.09*
Inflation	−0.06	−0.02	0.03	−0.04	0.11	0.06	0.13	−0.01
Std. error	*0.02*	*0.02*	*0.02*	*0.03*	*0.02*	*0.02*	*0.02*	*0.03*
GDP growth	5.48	8.68	8.38	2.82	4.54	1.89	2.67	2.53
Std. error	*1.38*	*1.32*	*1.34*	*1.37*	*1.49*	*1.37*	*1.50*	*1.38*
Trade/GDP	−0.22	1.94	2.17	−0.16	1.77	0.47	0.71	−5.13
Std. error	*1.70*	*1.73*	*1.61*	*1.78*	*1.38*	*1.58*	*1.67*	*1.64*
Real FX	−10.10	−7.91	−3.86	−5.44	−2.68	−10.01	−12.11	−10.92
Std. error	*0.48*	*0.59*	*0.79*	*0.43*	*0.62*	*0.67*	*0.77*	*0.71*

Regression of economic variables on a constant and two liberalization measures. The first is a transition indicator (= 1 for $-t-2$ to $t+2$) and second is the liberalization indicator (= 1 for $t+3$ to $t+62$). We report standard errors corrected for groupwise heteroskedasticity. Credit rating, GDP growth, and Trade/GDP are available only at semiannual, annual, and annual frequencies, respectively. Repeated monthly values are used in these cases and standard errors are inflated by the square root of the number of repetitions per observation to account for the repetition. Units are percent except for credit rating, correlation, beta, and real FX level.

minimum and 122.5 (76.9) for the five-year (three-year) regressions. Hence, our panels are unbalanced.

We estimate Eq. (10) using three different estimators: pooled OLS, groupwise heteroskedasticity correction, and groupwise heteroskedasticity plus AR(1) error term correction. We also considered both fixed-effects and a constant intercept. While all the results are available on the Internet, Table 6 only reports the results for the no-fixed effects case with groupwise heteroskedasticity correction. Since the model is in event time, we have $Trans_{t,i} = Trans_t$ and $Lib_{t,i} = Lib_t$; moreover, the constant is uncorrelated with the regressors. Hence, fixed effects are unnecessary and simply exhaust degrees of freedom. Pooled OLS will uncover the average change in means, which is what we are after, but the groupwise heteroskedasticity results are less sensitive to outlier observations. Most results are qualitatively robust across estimation methods.

We select eight different dates for the analysis. The first four dates are dates discussed and used in Bekaert and Harvey (2000). Three are purely exogenous dates: the introduction of American Depositary Receipts (ADRs), the introduction of country funds, and official liberalization dates. The fourth date uses the break point in the U.S. equity holdings based on the univariate analysis with breaks in all the parameters of the autoregression. The next four dates are endogenous break dates. We investigate the two bivariate breaks (returns and dividend yields and returns and market capitalization to GDP, both with the coefficients on the world instruments breaking), as well as the break dates for the trivariate and quadravariate systems. We discuss the five-year results but we note the differences from the three-year analysis below.

The first set of variables is directly related to the process of market integration. We expect the market integration process to lead to permanent price increases that decrease dividend yields and expected returns, conditional on foreigners indeed increasing their holdings of the local market capitalization (see Bekaert and Harvey, 2000; Bekaert et al. 2002). Increased integration and the global pricing it entails can also imply higher betas and correlations with respect to the world market return. Although stories about foreigners inducing excess volatility abound, it is not clear theoretically what should happen to local market volatility post-liberalization. Higher stock market prices can also lead to increased IPO activity and the associated stock market development can generally lead to larger market capitalizations.

The results in Table 6 are very much as expected. Average returns go down post-liberalization in five of eight cases, with the sharpest decrease occurring for the date from the quadravariate system (a significant 2.07% drop). For three dates, average returns actually increase and not all changes are significantly different from zero. However, returns are very noisy and dividend yields are likely a better indicator of permanent price changes: indeed, the decrease in dividend yield varies between 1.23% in the quadravariate system and the capital flow break to 2.98% for the country fund introduction break. Dividend yields always decrease and the decrease is always statistically significantly different from zero. There is only one country (Indonesia) for which dividend yields and returns both increase, which is inconsistent with a market integration story.

The other results also match up with expectations: market capitalization to GDP generally increases, in some cases by more than 10%, and U.S. holdings increase on average by 0.6 to 3.5%—the largest increase, not surprisingly, occurring for the capital flow break point. To put these values in perspective, the average level of U.S. holdings before the capital flow break point is 2.8%. The coefficient estimate in Table 6 implies a more than doubling of U.S. participation in these markets.

Consistent with the results in Bekaert and Harvey (2000), betas and correlations increase in a statistically significant way.[20] Again, these increases are economically important. For example, the average level of beta before the quadravariate break is 0.25. The results in Table 6 suggest that this beta increases on average by 0.18.

Table 6 also examines the impact on volatility. Confirming earlier results in De Santis and İmrohoroğlu (1997), Bekaert and Harvey (1997), and Kim and Singal (2000), the impact on volatility is ambiguous across different dates. For example, the trivariate system suggests a volatility decrease. However, two of the exogenous dates, official liberalizations and country fund introductions, suggest a volatility increase.

The second set of variables focuses on stock market development. Of course, the sizable increase in market capitalization to GDP is one indicator of increased stock market development. In addition, the significant increases in turnover and value traded to GDP suggest increased trading activity and liquidity post-break. The concentration ratio robustly decreases but the results for the cross-sectional standard deviation are mixed. If markets were relatively underdeveloped before the liberalization, we would expect the cross-sectional standard deviation to increase with increased stock market development, but if global capital flows drive price changes post-liberalization, the cross-sectional standard deviation might actually decrease.

The final category we examine is macroeconomic aggregates. Credit ratings increase significantly for all countries. This measure is correlated with the cost of capital (see Erb et al. 1996). We find little relation between inflation and the size of the trade sector and our dates. This is surprising in the light of the fact that trade and financial liberalization often occur simultaneously (see Henry, 2000a). GDP growth increases substantially post-liberalization, with the change in GDP growth varying between 1.89% and 8.38%. This growth result is stronger than the evidence presented in Bekaert et al. (2001) for a larger cross-section of countries. However, in Table 6, there are no control variables in the regression—the regression only uses data just before and after the liberalization and the standard errors are likely somewhat underestimated given the autocorrelation present in GDP growth rates. The results for the specification with AR(1) errors are then also somewhat weaker.

We also find that on average the real exchange rate robustly decreases, implying a local currency appreciation. Kim and Singal (2000) do not find real exchange rate appreciations after stock market openings. Our results seem more in line with the

[20] We also examined three-year rolling unconditional betas against the world index using both IFC global and investable indices. The results using these alternative betas are consistent with those reported in Table 6. For example, in the quadravariate system, the global beta increases by 0.27. Detailed results are available on the Internet.

conventional wisdom in international economics that capital inflows induce real exchange rate appreciations (see, e.g., Reisen, 1993). A real exchange rate appreciation can be the by-product of an exchange rate stabilization program, which manages to abruptly reduce the volatility of exchange rate changes, but where inflation converges more slowly to a lower level (see Reinhart and Vegh, 1995). Hence, macroeconomic reform accompanying financial liberalization might be the driving force behind this result. Nevertheless, there are no clear patterns in exchange rate volatility before or after liberalizations, although the exogenous break dates produce volatility decreases. Moreover, there are very large cross-country differences in the change in the real exchange rate. The number of countries with real rate depreciations varies between two and seven depending on the date chosen.

As mentioned earlier, we conduct two robustness experiments. First, we consider a three-year versus five-year before/after window. The results are largely robust to the change in window and are reported on the Internet. Second, we estimated the regression in Table 6 with six different econometric assumptions. We have argued that the most appropriate model is the model without fixed effects and correcting for groupwise heteroskedasticity. While the serial correlation correction makes a difference for the estimates in a number of cases, the broad analysis is resilient to the econometric assumptions.

Of course, our analysis now leaves us with eight different candidate break dates. Which date should we choose? Since we have argued that market integration is an all-encompassing structural break that should affect all economic variables, perhaps the answer is the date that yields the "strongest" breaks. It is nearly impossible to formalize the notion of "strongest break", but Table 7 makes an attempt in this direction. In the context of this paper, where we focus on changes in the means of relevant economic variables, the market integration date should be the date that leads to the sharpest changes in the variables of interest. The only meaningful way to aggregate across diverse variables is to aggregate statistically: that is, Table 7 essentially produces chi-square statistics by aggregating the squared *t*-statistics of the liberalization slopes of the regressions in Table 6. Of course, these variables are correlated, and we do not attempt to correct for this correlation, since our purpose is to provide a summary measure of impact rather than a statistical measure of integration. Clearly, the changes in most of our variables are very statistically significant. We are only interested in comparing the standardized magnitude of the changes for different economically meaningful variables across different dates.[21]

The first panel in Table 7 focuses on variables closely related to market integration. We use returns, dividend yields, and market capitalization to GDP as the basic variables, then separately add beta, correlation, and credit ratings. The first major result is that country fund introductions and official liberalizations generate the largest mean changes, with the country fund introductions coming in first. The importance of country funds has been noted before (see Errunza et al. 1999). These

[21] The sensitivity of Table 7 results to different econometric assumptions and before/after windows is presented on the Internet.

Table 7
Sweeping changes as measured by different break dates

	ADR introduction	Official liberalization	Country fund intro	Univariate: holdings/ mkt.cap.	Bivariate 1: returns and dividend yields	Bivariate 2: returns and mkt.cap./GDP	Trivariate	Quadravariate
Market integration								
I. Returns, dividend yields and holdings/ mkt.cap.	633.89	1009.10	1616.18	890.64	546.17	222.62	340.25	536.41
II. I + beta	1097.21	1388.16	1701.69	1532.26	614.54	913.38	644.51	966.03
III. I + correlation	864.70	1325.48	1696.76	1046.21	584.83	961.47	466.81	776.66
IV. I + credit rating	672.52	1029.25	1645.33	910.19	547.01	249.87	341.32	593.91
Stock market development								
Mkt.cap./GDP, turnover, and value traded/GDP	492.92	588.07	1164.20	419.29	425.19	1309.97	991.19	727.14
Macroeconomic variables								
I. GDP growth, trade/ GDP and inflation	22.20	45.34	43.72	5.93	34.56	10.48	44.62	13.47
II. I + credit rating	60.83	65.50	72.86	25.48	35.40	37.72	45.69	70.97
III. I + real FX rate	466.06	224.58	67.91	162.86	52.92	230.76	294.34	250.83

We construct aggregate measures of how "sweeping" the changes associated with each date are by summing the squared t-statistics implied by the coefficients and standard errors from Table 6.

results suggest that regulatory changes lead to important changes in many financial variables. Among the endogenous dates, the capital flow break date scores the best, closely followed by the quadravariate break date. The reason the endogenous dates perform worse than the exogenous dates should not be very surprising. We find that including all parameters relative to just the mean is important in determining statistically significant breaks. Clearly, the break dates have partially identified the changes in the dynamics of key financial variables, rather than just the means. Bekaert et al. (2002), for example, show very substantial differences in impulse responses of expected returns with respect to shocks in world interest rates, capital flows, and other variates, pre- and post-liberalization.

When we introduce betas and correlation, the bivariate market capitalization to GDP system generates relatively large changes. This system also produces the highest statistic for the stock market development variables. Hence, stock market development seems important in driving the correlation with world markets and it is not perfectly correlated with official liberalizations. Nevertheless, country fund introductions are also associated with large changes in market capitalization and value traded.

Finally, for macroeconomic variables the rankings across break dates are very sensitive to what variables are included. Interestingly, the capital flows date does not appear to be very important with respect to macroeconomic aggregates. This is even true when real exchange rates are considered, although real exchange rate appreciations are often ascribed to excessive foreign capital inflows.

6.2. *Integrated versus segmented portfolios*

To obtain an alternative view of the financial effects of market integration, we form equally weighted portfolios of countries before and after their break points. The pre-break point portfolio we call the ‘‘segmentation portfolio’’ and the post-break point portfolio we call the ‘‘integration portfolio’’. We do not include a country in either portfolio in the five months around the break point. We present summary statistics on these portfolio returns in Table 8. We focus on the January 1990–December 1993 period when many of the countries experience a break.

There are substantial differences among the returns of the integrated and segmented countries across the different break date choices but the changes are only significant in the case of official liberalizations, the bivariate market capitalization to GDP system, and the trivariate and quadravariate systems. The largest difference in returns (Panel A) is found with the bivariate (returns and market capitalization/GDP) as well as the quadravariate system. Average returns are sharply lower for the endogenous break dates in this group in the post-break period, but higher in the official liberalization case. Perhaps the transition period we specify is not long enough to fully exclude the return to integration. Note that returns are also much lower for countries with country fund introductions. Portfolio volatility is higher in half of the cases and often the change in volatility is small.

Panel B of Table 8 presents the correlations of the various portfolios. The average correlation of the integrated (segmented) portfolios is 0.73 (0.66). However, the

Table 8
Integration/segmentation portfolio analysis for January 1990–December 1993

A. Characteristics of portfolios

	ADR introduction (ADR)		Official liberalization (OL)		Country fund intro (CFI)		Univariate: holdings/GDP (UNI)	
	Segmented	Integrated	Segmented	Integrated	Segmented	Integrated	Segmented	Integrated
Mean	32.03	27.47	6.58	34.73	32.77	20.47	25.51	24.15
Std. dev	17.02	24.75	18.59	17.31	22.22	18.94	13.99	27.04
Corr.	0.38	0.27	0.35	0.34	0.12	0.47	0.42	0.34
Beta	0.71	0.73	0.72	0.64	0.30	0.98	0.65	1.03
t-stat (S = I)	−0.36		2.51		−0.86		−0.12	

	Bivariate 1: returns and div. yields (BIV1)		Bivariate 2: returns and mkt.cap./GDP (BIV2)		Trivariate (TRI)		Quadravariate (QUAD)	
	Segmented	Integrated	Segmented	Integrated	Segmented	Integrated	Segmented	Integrated
Mean	15.27	25.59	40.67	−1.25	36.42	14.73	32.68	7.63
Std. dev	17.15	16.14	17.41	20.92	18.48	18.84	14.64	24.10
Corr.	0.16	0.46	0.32	0.50	0.20	0.54	0.26	0.52
Beta	0.30	0.83	0.62	1.17	0.42	1.12	0.42	1.39
t-stat (S = I)	−0.02		−3.97		−2.10		−2.17	

B. Correlations of portfolios

		Integrated								Segmented							
		ADR	OL	CFI	UNI	BIV1	BIV2	TRI	QUAD	ADR	OL	CFI	UNI	BIV1	BIV2	TRI	QUAD
Integrated	ADR	—				Average for integrated								Average of integrated/segmented by break (diagonal)			
	OL	**0.76**	—			0.73								0.34			
	CFI	**0.68**	**0.85**	—													
	UNI	**0.82**	**0.80**	**0.81**	—												
	BIV1	**0.65**	**0.89**	**0.91**	**0.76**	—											
	BIV2	**0.40**	**0.64**	**0.76**	**0.55**	**0.72**	—										
	TRI	**0.47**	**0.74**	**0.87**	**0.56**	**0.83**	**0.83**	—									
	QUAD	**0.42**	**0.71**	**0.88**	**0.62**	**0.81**	**0.86**	**0.92**	—								
Segmented	ADR	**0.31**	0.66	0.60	0.45	0.75	0.47	0.60	0.51	—				Average for segmented			
	OL	0.19	**0.22**	0.35	0.22	0.49	0.31	0.43	0.32	**0.68**	—			0.66			
	CFI	0.15	0.37	**0.04**	0.10	0.36	0.10	0.12	0.03	**0.68**	**0.50**	—					
	UNI	0.54	0.87	0.83	**0.58**	0.93	0.68	0.80	0.74	**0.84**	**0.53**	**0.54**	—				
	BIV1	0.40	0.67	0.46	0.42	**0.45**	0.32	0.35	0.32	**0.60**	**0.20**	**0.53**	**0.68**	—			
	BIV2	0.57	0.81	0.74	0.62	0.85	**0.40**	0.63	0.55	**0.85**	**0.52**	**0.53**	**0.88**	**0.63**	—		
	TRI	0.54	0.78	0.62	0.63	0.75	0.39	**0.38**	0.38	**0.76**	**0.31**	**0.60**	**0.79**	**0.67**	**0.83**	—	
	QUAD	0.56	0.79	0.63	0.58	0.79	0.40	0.51	**0.37**	**0.85**	**0.48**	**0.66**	**0.85**	**0.67**	**0.89**	**0.91**	—
	World	0.27	0.34	0.47	0.34	0.46	0.50	0.54	0.52	0.38	0.35	0.12	0.42	0.16	0.32	0.20	0.26

Portfolios are constructed over the specified time period. From the sample of 20 countries, the "Segmented" (S) portfolio includes countries whose liberalization is at least three months away (in the future). Similarly, the "Integrated" (I) portfolio includes countries whose liberalization occurred at least three months ago. Countries within two months of their liberalization (in either direction) are not in either portfolio. Portfolios are equally weighted and rebalanced monthly. A portfolio must have at least two component returns to have a return counted for that period. This only matters for the "segmented univariate" portfolio, which has no observations after March 1993. Units are annualized percent. t-stat (S = I) is a two-sided t-test of whether the mean returns are different from segmented to integrated. World represents correlation with world market return.

correlation of the integrated and segmented portfolios across the eight different break definitions is only 0.34. Based on endogenous break dates, the integrated portfolio most correlated with both the official liberalization break date portfolio and the country fund introduction break date portfolio is the one using the bivariate returns–dividend yield system (BIV1 in the table). Among other endogenous break point portfolios, BIV1 is the most correlated with the portfolio based on U.S. market holdings. The latter portfolio is, of all endogenous break point portfolios, most highly correlated with the one based on ADR introductions. Both ADR introductions and capital flow break points tend to occur later in the sample.

Two measures of the comovement of the portfolios and the world market portfolio, correlation and beta, are also examined in Panel A of Table 8. In Table 6, betas and correlations increase across all different datings. Here, this only happens when using the four multivariate endogenous break dates or country fund introduction dates to define integration. The correlation with the world market usually more than doubles moving from the segmented to the integrated portfolio. The beta also shows a dramatic increase—also more than doubling in most of the samples. These results are consistent with Bekaert and Harvey (2000).

It is important to note that the increases in the comovement measures do not imply that the diversification benefit of investing in emerging markets disappears. For example, in the quadravariate system, the correlation increases from 0.26 to 0.52. This is still less than the correlation among large developed markets. Over the same time period, the average correlation of France, Germany, Italy, and U.K. returns with the world return is 0.65. These results indicate that the emerging market returns are much more sensitive to world factors post-integration. There are two potential sources for this increased correlation. In a world where discount rates vary through time, global pricing can induce higher correlations with the world market as opposed to local pricing. It is also possible that the local firms cash flows become more sensitive to world factors. One potential channel is the increased trade flows occurring after financial and economic integration.[22]

6.3. *When does integration occur?*

There are eight potential break dates—but when does integration occur? The dates in Table 3 include exogenous dates (which include official liberalizations, introduction of the first U.S. ADR, and introduction of country funds) as well as endogenous dates from the break analysis.

There are many different approaches to determining when market integration occurs. One strategy is to date the liberalization by the official government decrees and assume that the liberalization date is the market integration date. The disadvantage of this approach is that investors can sometimes bypass regulations. Alternatively, investors might not place much faith in the longevity of the new

[22] For example, based on the bivariate returns and dividend yield system, 16 of 17 countries experience an increase in the ratio of trade to GDP after the break point.

regulations. The approach in our paper is to let the data tell the story. However, it is also important to link our break dates to what is happening in each market.

One country we have focused on is Colombia. Our chronology of Columbia's history in Section 2.2 shows 1991 being a critical year for capital market reforms. In October 1991, the peso is deregulated, foreign firms are allowed to remit 100% of their profits, and the broad reforms of Resolution 51 take effect. In December 1991, Resolution 52 allows foreigners to purchase up to 100% of locally listed companies. The median break date in the bivariate analysis of dividends and returns is October 1991. This is also the date chosen by the trivariate analysis. However, the quadravariate analysis chooses a later date, May 1994, and is heavily influenced by the increase in net equity capital flows that occurs in early 1994.

There are many countries with a close relation between the reforms and the break dates. For example, the bivariate (returns and dividend yields) break date for Argentina is February 1989. In March 1989, the Brady plan (an adjustment package that combined debt relief and market-oriented reforms) is announced. By November 1989, the New Foreign Investment Regime is put into place. All legal limits on foreign investment are abolished. Capital gains and dividends can be repatriated with no need for previous approval of transactions. Legal limits regarding the type or nature of foreign investment are abolished. A free exchange regime (free repatriation of capital, remittance of dividends and capital gains) is introduced.

Similar to the experience with Colombia, the addition of market capitalization to GDP and holdings to GDP pushes the Argentinian break dates forward in time. Indeed, a closer examination suggests that the New Foreign Investment Regime is not fully implemented until later. Capital controls are not formally abolished until March 1991. It is only in August 1991 that a law is enacted to protect U.S. dollar accounts. Indeed, a new currency is introduced in January 1992. Given this activity and perhaps some doubts about the credibility and longevity of the reforms, foreign capital flows do not increase until well after the reforms of November 1989. Our quadravariate analysis suggests a break in June 1992.

Finally, consider the case of Turkey. In 1988, Article 15 of Decree 32 paves the way for foreigners to invest directly in Turkey. As a result, foreigners are no longer required to seek preapproval to purchase or sell securities listed on the Istanbul Stock Exchange. Our trivariate and multivariate break analyses suggest that the break occurs in September 1989 and May 1989, respectively. These endogenous dates are very close to the exogenous dates.

Table 9 analyzes the association between the endogenous and exogenous dates. For each country, we calculate the raw number of months between these break dates. We add the univariate returns break date to our list of endogenous dates. Panel A presents the average of the absolute differences and Panel B reports the average difference (endogenous minus exogenous). There are two insights from this analysis. First, adding additional information (multivariate systems) makes the endogenous break dates closer to the exogenous dates. The two lowest average absolute differences are the trivariate and quadravariate systems. Second, the average difference suggests that the endogenous break dates are later than the exogenous

Table 9
Association between endogenous and exogenous break dates

	Official liberalization	ADR introduction	Country fund introduction	Average
Panel A: Average absolute difference (months) between endogenous and exogenous break dates				
Univariate 1: returns	46.9	51.6	45.4	48.0
Univariate 1: holdings/mkt.cap.	36.8	26.8	59.5	41.0
Bivariate 1: returns and div. yields	45.7	37.9	30.1	37.9
Bivariate 2: returns and mkt.cap./GDP	35.0	21.9	47.5	34.8
Trivariate	33.2	30.9	35.1	33.1
Quadravariate	33.3	23.2	44.6	33.7
Panel B: Average difference (months) between endogenous and exogenous break dates				
Univariate 1: returns	−35.6	−46.2	−7.2	−29.7
Univariate 1: holdings/mkt.cap.	33.4	18.8	56.8	36.3
Bivariate 1: returns and div. yields	−26.3	−31.6	3.6	−18.1
Bivariate 2: returns and mkt.cap./GDP	12.9	8.4	47.2	22.8
Trivariate	−3.4	−11.9	23.7	2.8
Quadravariate	21.9	7.2	44.6	24.6

For each country, we calculate the raw number of months between various endogenous and exogenous break dates (the convention here is endogenous minus exogenous). The exogenous dates (official liberalization, ADR introduction, and country fund introduction) are from Bekaert and Harvey (2000). The univariate returns break is from Table 2 (all coefficients breaking with world information). All other break dates are found in Table 3. In Panel A, the absolute value of this number is calculated and then averaged across countries. In Panel B, the raw difference is averaged.

dates. The exceptions are the bivariate returns/dividend yield and the univariate returns systems. The dividend yield date is most closely associated with country fund introductions. Hence, dividend yield breaks precede official liberalizations which, in turn, precede break dates where the market capitalization or capital flow variables play a role.

There are a number of potential interpretations of these results. The late break dates are likely consistent with the notion that market integration is a gradual process. Many foreign investors may be skeptical of official pronouncements, and foreign capital flows and large changes in market capitalization occur much later than official liberalizations. Alternatively, it is striking that the market capitalization dates seem to be clustered in 1993. This is the year when capital flows to emerging markets really accelerate. In April 1993, the U.S. interest rate reaches a trough and some believe that this triggered capital outflows from the U.S. to emerging (and other) markets. The market capitalization break dates are in fact much more closely associated with the April 1993 date (an average of only 21 months) than with the capital flow dates. Of course, 1993 also witnessed numerous new ADR programs, sometimes coupled with large privatization programs which may have increased local market capitalization. Overall, the endogenous break dates are more closely aligned with ADR launchings than with official liberalizations or country fund introductions.

7. Conclusions

We study structural breaks in a host of financial and economic time series in 20 emerging markets. The methodology, developed in Bai et al. (1998), allows the determination of the break date with a 90% confidence interval. In contrast to previous empirical applications, we allow for all of the parameters in the regression to change at the break point. If the recent capital market liberalization process in emerging markets effectively integrated these markets into world capital markets, we expect the move from segmentation to integration to be accompanied by a significant break in a number of time series.

We find strong evidence of structural breaks in emerging equity markets. The statistical significance of our results, and thus the precision of our dating exercise, is enhanced by allowing all parameters to break (rather than only the mean), by examining variables such as dividend yields which capture permanent price level changes (rather than only noisy returns data), and by examining multiple time series simultaneously. The confidence interval around our break dates is surprisingly tight given that the liberalization process is often complex and gradual. We find that integration brings about or is accompanied by an equity market that is significantly larger and more liquid than before, and stock returns that are more volatile and more correlated with world market returns than before. Integration is also associated with a lower cost of capital, an improved credit rating, a real exchange rate appreciation, and increased real economic growth. These results are based on cross-sectional averages and the dispersion in the changes in financial and economic time series across countries may be wide. We find no evidence of significant structural breaks in developed markets.

Our dates should be useful for the rapidly growing body of literature studying the changes in emerging markets after liberalization. Consider three possible alternative measures of a break date: a date based on major regulatory reforms liberalizing foreign equity investments, the date of the announcement of the first ADR issue, and the date of the first country fund launching. Our endogenous structural break dates are mostly within three years of one of these dates, but the timeliness and identity of the closest exogenous break date varies greatly across countries. Generally, though, endogenous dates occur later than exogenous break dates and are most closely associated with ADR introductions. Allowing foreign investment does not appear to be sufficient to bring about market integration; foreigners still have to be willing to invest.

Appendix

The critical values for dimensions up to 68 are given in Table 10.

Table 10
Asymptotic critical values for $\max_k F(k)$ test statistic: tests for a simultaneous break in q parameters

Dimension q	15%	10%	5%	1%
1	6.343	7.154	8.692	11.809
2	9.056	10.226	11.780	16.366
3	11.361	12.460	14.331	17.870
4	13.135	14.030	15.655	19.384
5	15.095	16.564	18.441	23.057
6	16.844	18.451	20.418	24.682
7	18.311	19.524	22.211	27.023
8	20.194	21.926	24.319	28.891
9	21.447	23.122	25.424	30.849
10	23.328	24.648	27.521	32.405
11	24.617	25.961	28.356	33.742
12	26.396	27.943	30.396	36.182
13	27.851	29.653	32.134	36.826
14	29.435	31.180	33.377	39.064
15	30.661	32.516	34.875	39.629
16	32.331	33.931	36.623	41.642
17	33.490	35.505	38.584	44.292
18	35.301	37.038	40.108	45.269
19	36.445	38.528	41.759	47.978
20	37.944	39.602	42.820	47.749
21	39.631	41.391	44.194	50.408
22	40.611	42.119	44.811	50.398
23	42.021	43.872	46.863	52.211
24	43.127	45.106	48.488	55.504
25	44.985	46.821	49.679	56.871
26	46.407	48.409	51.352	56.929
27	47.333	49.441	52.126	58.901
28	48.522	50.450	53.548	61.365
29	50.110	51.934	55.465	61.648
30	51.668	53.392	56.989	63.824
31	53.090	55.766	59.522	65.642
32	53.442	55.686	59.504	66.638
33	55.304	58.014	61.440	67.576
34	56.427	58.327	61.420	67.668
35	57.908	60.384	63.753	69.472
36	59.360	61.432	64.393	72.986
37	60.621	63.133	66.590	74.534
38	61.090	63.336	67.080	75.309
39	62.716	65.295	69.711	76.520
40	64.718	67.330	70.768	76.730
41	65.574	67.924	71.617	78.904
42	66.447	68.860	72.265	80.179
43	67.870	70.255	74.240	81.535
44	69.830	71.740	75.748	83.884
45	70.403	72.734	75.834	84.306
46	71.979	74.340	78.647	86.463
47	73.166	75.481	79.340	87.366
48	74.323	76.852	80.354	88.284
49	75.769	78.206	83.066	89.699

Table 10 (*continued*)

Dimension q	15%	10%	5%	1%
50	77.509	79.589	83.502	91.599
51	77.569	80.698	84.991	93.892
52	78.909	81.712	85.904	93.606
53	80.716	83.268	87.374	97.320
54	81.683	84.416	88.305	97.371
55	83.256	85.649	88.891	97.624
56	84.569	86.806	91.200	98.058
57	85.899	88.408	93.138	100.505
58	86.255	88.910	92.911	100.552
59	88.730	91.093	95.205	102.933
60	89.950	93.104	97.295	106.150
61	90.519	93.187	97.068	107.428
62	91.145	94.197	99.084	108.896
63	92.377	95.045	99.259	108.092
64	94.331	96.567	100.624	112.716
65	95.534	98.123	102.389	111.575
66	96.531	99.435	103.219	111.646
67	97.780	100.962	105.699	113.928
68	98.514	101.635	106.481	115.907

q represents the dimension of the test statistic, $F_T(k)$ (the dimension of $S\delta$ in (6)). 15% trimming is used. Values in this table are based on 2000 Monte Carlo replications of a discrete approximation to the limiting representation (F^*) of these statistics as functionals of a q-dimensional Brownian motion, with a discretization grid of 5000. See Section 3.3 in the text for details.

References

Aggarwal, R., Inclan, C., Leal, R., 1999. Volatility in emerging stock markets. Journal of Financial and Quantitative Analysis 34, 33–55.

Alexander, G.J., Eun, C.S., Janakiramanan, S., 1987. Asset pricing and dual listing on foreign capital markets: a note. Journal of Finance 42, 151–158.

Bai, J., Perron, P., 1998. Estimating and testing linear models with multiple structural changes. Econometrica 66, 47–78.

Bai, J., Lumsdaine, R.L., Stock, J.H., 1998. Testing for and dating breaks in stationary and nonstationary multivariate time series. Review of Economic Studies 65, 395–432.

Banerjee, A., Lumsdaine, R.L., Stock, J.H., 1992. Recursive and sequential tests of the unit root and trend-break hypotheses: theory and international evidence. Journal of Business and Economic Statistics 10, 271–287.

Bekaert, G., Harvey, C.R., 1995. Time-varying world market integration. Journal of Finance 50, 403–444.

Bekaert, G., Harvey, C.R., 1997. Emerging equity market volatility. Journal of Financial Economics 43, 29–78.

Bekaert, G., Harvey, C.R., 1998. Capital flows and the behavior of emerging market equity returns. Unpublished Working Paper 6669. National Bureau of Economic Research, Cambridge, MA.

Bekaert, G., Harvey, C.R., 2000. Foreign speculators and emerging equity markets. Journal of Finance 55, 565–613.

Bekaert, G., Harvey, C.R., Lundblad, C.T., 2001a. Emerging equity markets and economic development. Journal of Development Economics 66, 465–504.

Bekaert, G., Harvey, C.R., Lundblad, C.T., 2001b. Does financial liberalization spur growth? Unpublished Working Paper 8245, National Bureau of Economic Research, Cambridge, MA.

Bekaert, G., Harvey, C.R., Lumsdaine, R.L., 2002. The dynamics of emerging market equity flows. Journal of International Money and Finance (21)3, 295–350.
Ben-David, D., Papell, D.H., 1998. Slowdowns and meltdowns—post-war growth: evidence from 74 countries. Review of Economics and Statistics 80, 561–571.
Bessimbinder, H., Chan, K., Seguin, P., 1996. An empirical examination of information, differences of opinion, and trading activity. Journal of Financial Economics 40, 105–134.
Bossaerts, P., Hillion, P., 1999. Implementing statistical criteria to select return forecasting models: what do we learn? Review of Financial Studies 12, 405–428.
Christie, W.G., Huang, R.D., 1995. Following the pied piper: Do indvidual returns herd around the market? Financial Analysts Journal 51, 31–37.
Cole, H., Obstfeld, M., 1992. Commodity trade and international risk sharing: How much do financial markets matter? Journal of Monetary Economics 28, 3–24.
Connolly, R.A., Stivers, C.T., 2001. Characterizing the intertemporal market-to-firm volatility relation: new time-series and cross-sectional evidence. Unpublished working paper, University of Georgia, Athens, GA.
De Santis, G., İmrohoroğlu, S., 1997. Stock returns and volatility in emerging financial markets. Journal of International Money and Finance 16, 561–579.
Devereux, M.B., Smith, G.W., 1994. International risk sharing and economic growth. International Economic Review 35, 535–551.
Domowitz, I., Glen, J., Madhavan, A., 1997. Market segmentation and stock prices: evidence from an emerging market. Journal of Finance 52, 1059–1086.
Erb, C.B., Harvey, C.R., Viskanta, T.E., 1996. Expected returns and volatility in 135 countries. Journal of Portfolio Management 22, 46–58.
Errunza, V.R., Losq, E., 1985. International asset pricing under mild segmentation: Theory and test. Journal of Finance 40, 105–124.
Errunza, V.R., Losq, E., Padmandahan, P., 1992. Tests of integration, mild segmentation and segmentation hypotheses. Journal of Banking and Finance 16, 949–972.
Errunza, V.R., Hogan, K., Hung, M.W., 1999. Can the gains from international diversification be achieved without trading abroad? Journal of Finance 54, 2075–2107.
Eun, C.S., Janakiramanan, S., 1986. A model of international asset pricing with a constraint on the foreign equity ownership. Journal of Finance 41, 897–914.
Foerster, S.R., Karolyi, G.A., 1999. The effects of market segmentation and illiquidity on asset prices: evidence from foreign stocks listing in the U.S. Journal of Finance 54, 981–1013.
Garcia, R., Ghysels, E., 1998. Structural change and asset pricing in emerging markets. Journal of International Money and Finance 17, 455–473.
Henry, P.B., 2000a. Equity prices, stock market liberalization, and investment. Journal of Financial Economics 58, 301–334.
Henry, P.B., 2000b. Stock market liberalization, economic reform, and emerging market equity prices. Journal of Finance 55, 529–564.
Jorion, P., Schwartz, E., 1986. Integration versus segmentation in the Canadian stock market. Journal of Finance 41, 603–613.
Kang, J.-K., Stulz, R.M., 1997. Why is there a home bias? An analysis of foreign portfolio equity ownership in Japan. Journal of Financial Economics 46, 3–28.
Kim, E.H., Singal, V., 2000. Opening up of stock markets: lessons from emerging economies. Journal of Business 73, 25–66.
Levine, R., Zervos, S., 1996. Stock market development and long-run growth. World Bank Economic Review 10, 323–340.
Lewis, K.K., 1996. What can explain the apparent lack of international consumption risk-sharing? Journal of Political Economy 104, 267–297.
Miller, D.P., 1999. The impact of international market segmentation on securities prices: evidence from Depositary Receipts. Journal of Financial Economics 51, 103–123.
Obstfeld, M., 1994. Risk taking, global diversification and growth. American Economic Review 84, 1310–1329.

Picard, D., 1985. Testing and estimating change-points in time series. Advances in Applied Probability 176, 841–867.
Reinhart, C.M., Vegh, C.A., 1995. Do exchange rates-based stabilizations carry the seeds of their own destruction? Unpublished working paper, International Monetary Fund, Washington, DC.
Reisen, H., 1993. The case for sterilized intervention in Latin America. Unpublished working paper. OECD, Paris.
Richards, A.J., 1996. Volatility and predictability in national markets: how do emerging and mature markets differ? Staff paper, International Monetary Fund, Washington, DC.
Schwarz, G., 1978. Estimating the dimension of a model. Annals of Statistics 6, 461–464.
Solnik, B., 1974. The international pricing of risk: an empirical investigation of the world capital market structure. Journal of Finance 29, 48–54.
Stapleton, R., Subrahmanyam, M., 1977. A multiperiod equilibrium asset pricing model. Econometrica 5, 1077–1093.
Stehle, R., 1977. An empirical test of the alternative hypothesis of national and international pricing of risky assets. Journal of Finance 32, 493–502.
Stulz, R.M., 1981. On the effects of barriers to international investment. Journal of Finance 36, 923–934.
Stulz, R.M., 1999. International portfolio flows and security markets. In: Feldstein, M. (Ed.), International Capital Flows. National Bureau of Economic Research, Cambridge, MA, pp. 257–293.
Tesar, L.L., Werner, I., 1995. U.S. equity investment in emerging stock markets. World Bank Economic Review 9, 109–130.
van Wincoop, E., 1994. Welfare gains from international risk sharing. Journal of Monetary Economics 34, 175–200.
World Bank, 1997. Private Capital Flows to Developing Countries: The Road to Financial Integration. National Bureau of Economic Research and Oxford University Press, Cambridge, MA.
Zivot, E., Andrews, D., 1992. Further evidence on the great crash, the oil price shock, and the unit root hypothesis. Journal of Business and Economic Statistics 10, 251–270.