School of Contemporary Chinese Studies

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UNITED KINGDOM · CHINA · MALAYSIA

WORKING PAPER SERIES

Stock market integration and financial crises: Evidence from Chinese sector portfolios

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Working Paper No. 14

January 2014

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Abstract

This paper assesses China's stock market integration with the global market during 1997-2013 and the impact of financial crises on the extent of integration within an augmented CAPM framework. We obtain time-varying global and national systematic risks for ten Chinese sectors from a state-space representation and investigate how these risks are priced in crisis and non-crisis periods, respectively, by employing a dummy variable approach and the Markov two-state regime-switching technique. We find evidence of positive exposures to both the global and national systematic risks in the majority of the ten sectors under study in the period excluding the recent global financial crisis, pointing to a conclusion of partial integration. During the global financial crisis, international influence on China's asset pricing strengthens, leading to complete market integration in half of the sectors under study. Complete integration is indeed a dominant feature of the global financial crisis as it is not evident in the high-volatility crisis periods. Our results suggest an opportunity of international diversification by selective inclusion of Chinese assets.

JEL classification: C32, F36, G12, G15

Key words: stock market integration; financial crisis; time-varying systematic risk; asset pricing; Kalman smoothing; regime switching;

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1. Introduction

It is widely accepted that the issue of whether stock markets are integrated or segmented has implications for financial decisions. Numerous studies have examined whether selected local markets are integrated with the global market. The study of stock market integration leads naturally to the question of whether such integration has reduced the benefits of international risk diversification and to a consideration of whether the answer to this question might alter during financial crises. Despite the considerable volume of research addressing these issues, few studies have concentrated on the case of China. The rapidly increasing importance of China within the global economy makes this a significant omission. This present study attempts to rectify the omission by investigating the impact of financial crises on China's asset pricing within the framework of an augmented CAPM.

We will investigate whether China's stock market is integrated with the global equity market and whether financial crises impact upon the extent of integration, using data for ten industry-specific sectors during 1997-2013. Following the established literature, e.g., Stehle (1977) and Jorion and Schwartz (1986), the empirical analyses will be carried out within the augmented CAPM framework, specifically an international CAPM augmented with orthogonalised national excess return and a domestic CAPM augmented with orthogonalised global excess return. We apply a Kalman smoothing approach to allow the parameters in the augmented CAPMs to vary over time. This is a well-established approach to permitting time-varying behaviour of systematic risk in a CAPM (see Ang and Chen, 2007; Mergner and Bulla, 2008; Li, 2013). The estimated time-varying global and national systematic risks are then employed within a two-beta asset pricing model in the style of Jorion and Schwartz (1986) to test for integration versus segmentation.

The two-beta asset pricing model can be regarded as a benchmark for how the Chinese sector portfolios should be expected to be priced, based on existing fundamentals. The pricing of the sector portfolios might potentially vary between tranquil and crisis periods because of changes in the exposures to global and national systematic risks. We will investigate how crises, especially the recent global financial crisis, have affected the exposures of the Chinese sectors to the systematic risks using the dummy variable approach and the Markov regime-switching approach respectively. Specifically, we will first introduce a dummy variable to indicate the global financial crisis period of August 2007-March 2009[1] and observe whether the interactions of the risk factor exposures with this dummy variable are statistically significant. Secondly, we will use the Markov two-state regime-switching technique (Hamilton, 1989 and 1990) to identify endogenously all the high-volatility crises that the Chinese sectors may have experienced during the period of January 1997-March

[1] We define the starting point of the global financial crisis as August 2007, when the major equity markets initially fell and central banks started intervening to provide liquidity to financial markets.

2013 and to produce separate estimates for the risk factor exposures during turbulent and tranquil periods.

We contribute to the literature in a number of important ways. Firstly, our study employs augmented CAPMs as an appropriate theoretical framework for assessing market integration and segmentation. Through investigating asset pricing at the sectoral level, we accommodate a more nuanced assessment of the extent of China's stock market integration with the global market, and of the consequent potential for risk reduction via international diversification. On the contrary, most previous studies have used statistical cointegration of national stock market indices or correlation between returns of national stock markets as evidence of market integration. Girardin and Liu (2007) argue that the existence of a long-run relationship between stock markets, identified through cointegration analysis, may imply cointegration of fundamentals, rather than integration of markets. Furthermore, Bekaert and Harvey (1995) question whether a correlation between the returns of national and world markets can be used to measure stock market integration. They argue that a country's stock markets could be perfectly integrated with the world market but still show a low or negative correlation because its industry mix differs significantly from the average world mix.

Secondly, we extend the augmented CAPM of Jorion and Scwartz (1986) to a time-varying setting using the Kalman filtering and Markov regime-switching techniques. Hence, we address the concern that the systematic risk is time-variant, solve the heteroskedasticity problem raised by Forbs and Rigobon (2002) and overcome the difficulty of arbitrarily dividing the sample into crisis and non-crisis periods. Thirdly, we compare asset pricing additionally during the crisis and non-crisis episodes within 1997-2013, in contrast to the majority of other studies which focus on a specific crisis, allowing for a generalisation of stock market integration or lack of it during times of stability and turbulence.

Finally, our findings have important implications for financial decision making and academic discussion. We find evidence of partial integration in the majority of the Chinese sectors in the period excluding the recent global financial crisis, suggesting potential for international risk diversification in China. In the global financial crisis, complete market integration becomes evident for more sectors, consistent with the literature which argues that stock market integration strengthens during crises. Although there is no evidence of complete integration using the regime-switching approach, international influence on China's asset pricing strengthens in the form of partial integration in a greater number of sectors in the high-volatility periods relative to the low-volatility periods. As far as we are aware, the possibilities that market integration may differ across sectoral portfolios and that the CAPM framework may be more or less applicable under different volatility conditions have received little attention hitherto. Our study therefore enriches the empirical literature on stock market integration.

The remainder of the paper is organised as follows. In section 2 we review the literature and in section 3 we describe the empirical framework. We carry out the preliminary data analysis in section 4 and implement estimation and tests in section 5. Section 6 concludes.

2. Review of the literature

Our work relates closely to the growing literature on stock market integration in the context of asset pricing models. In an asset pricing sense, integration is defined as a situation where investors earn the same risk-adjusted expected return on similar financial instruments in local and global markets. Hence the investigation of market integration for a particular economy could, in principle, involve testing whether systematic risk relative to the global market remains the only significant factor in an international capital asset pricing model when this has been augmented to include a national systematic risk. Discovering that the national risk makes a significant contribution additional to that from the global risk is evidence against the integration hypothesis. Conversely, discovering that only the national risk is significant in a national/global two-beta model would be evidence in favour of segmentation.

Using the traditional two-pass approach to parameter estimation, Stehle (1977) finds that the pricing of US securities is significantly related to a global market portfolio. Using a maximum likelihood approach to estimate all parameters simultaneously, Jorion and Schwartz (1986) find strong evidence of segmentation relative to the North American market in the pricing of Canadian stocks during 1963-1982. Mittoo (1992) re-examines the integration between the Canadian and US stock markets during 1977-1986 using the augmented CAPM and a multi-factor pricing model. Both models suggest that segmentation occurs during 1977-1981 and integration during 1982-1986. Wang and Di Iorio (2007) apply the approach of Jorion and Schwartz (1986) to investigate the segmentation / integration of three China-related stock markets and find that Chinese A-, B- and H-share markets were all segmented from the global market during 1995-2004.

In response to the studies, such as Bos and Newbold (1984) and Collins *et al.* (1987), which suggest that systematic risks relative to national and global markets are not constant over time, models with time-varying conditional covariance risk have been used to investigate integration between national markets. For example, in the framework of a multivariate GARCH model, De Santis and Imrohoroglu (1997) derive time-varying conditional covariances between some national indices and the global index and assess integration by its testable implication that the price of covariance risk is positive and equal across the national markets under study. More recent studies, for example Hardouvelis *et al.* (2006), follow Bekaert and Harvey (1995) and include the time-varying variance of a national index to represent the domestic factor. By applying time-varying weights to the prices of the covariance and variance, they estimate the time-varying probability that national stock markets are segmented or integrated. However, Ang and Chen (2007) argue that the time-varying conditional covariance and variance derived from the GARCH models are strictly driven by past innovations in returns and do not have independent random components. Li (2013) uses the Kalman smoothing technique to derive time-varying national and systematic risks of augmented CAPMs and test for integration versus segmentation within the framework of a two-beta CAPM in the style of Jorion and Schwartz (1986). She finds evidence of partial integration in terms of positively priced national and global systematic risks for the Chinese A- and B-share market indices during 2000-2010.

A question of particular interest for international investors and policymakers is whether stock market integration strengthens in times of crisis. Longin and Solnik (1995) document that the interdependence of national equity markets increases over time due to growing international economic, political and financial integration and the common factors that affect all economies simultaneously. Subsequent studies, such as Karolyi and Stulz (1996) and Ang and Chen (2002), confirm that markets are more highly correlated in periods when global factors, such as crises, cause stock market turbulence. Similarly, Frankel and Schmukler (1998) find evidence of stronger correlations in returns between the Latin American emerging markets during the 1994 Mexican crisis while Baig and Goldfajn (1999) tentatively suggest an increase in correlation between the Asian stock markets during the 1997 East Asian crises.

However, correcting for the change in the volatility of returns in the country where the crisis originates, Forbes and Rigobon (2002) find that correlation coefficients across multi-country returns are not significantly higher during crisis periods such as the 1987 US stock market crash, the 1994 Mexican peso crisis and the 1997 East Asian crisis. More recent studies, such as Corsetti *et al.* (2005) and Chiang *et al.* (2007), have questioned the methodology of Forbes and Rigobon (2002) and suggest that the omission of the country-specific component of the change in the volatility of returns or a dynamic increment in correlation has led to an under-estimation of the cross-market correlation during crisis periods in Forbes and Rigobon (2002). As a result, some have used alternative techniques to overcome the limitations associated with correlation analysis. For example, while Pelletier (2006) uses a regime-switching model in the context of correlation, Boyer *et al.* (2006) use both a regime-switching model and extreme value theory for numerous accessible and inaccessible stock indices, suggesting that crises spread through the asset holdings of international investors rather than through changes in fundamentals. Chiang *et al.* (2007) apply the dynamic conditional correlation (DCC) from the multivariate GARCH to nine Asian stock markets over 1990-2003 and find an increase in correlation during the Asian-crisis period. Rodriquez (2007) models stock market dependence with regime-switching copulas and finds evidence of changing dependence and asymmetry during the Asian and the Mexican crises. These recent studies have suggested, as summarised by Beine *et al.* (2010), that in times of financial crises, stock market integration increases because fundamentals deteriorate more or less simultaneously across countries. Hence stronger integration with the global market during crises may not result in the expected effects on international risk diversification or economic growth.

Although various capital asset pricing models have been developed to address the question of market integration, very few studies in this line of research have investigated the impact of financial crises on stock market integration. Bekaert *et al.* (2005) apply a two-factor model with time-varying betas and measure the impact of financial crises as correlation among residuals. Based on the latent factor model, Bekaert *et al.* (2005) decompose the market correlations into market interdependence and contagion. They find that the Mexican crisis contributes to increasing market interdependence while the East-Asian crisis leads to both stronger market interdependence and contagion across markets. Our work owes most to the asset pricing approach in Bekaert *et al.* (2005), testing integration in a time-varying setting. What we add is the separate estimation of market interdependence in times of stability and turbulence through a regime-switching technique.

3. Empirical framework

3.1. The two-beta asset pricing model

According to modern portfolio theory, in fully integrated markets investors are indifferent between the national markets to which they allocate funds because there is a common reward per unit of systematic risk exposure. In this case, systematic risk relative to the global market should be the only determinant in asset pricing. Where (say) capital controls, differential tax treatments or informational costs are imposed on some stock markets, then these markets may be segmented relative to global markets. Full segmentation implies that only the domestic factors, reflecting systematic risk relative to the national market, should enter the pricing of assets. Stehle (1977) and Jorion and Schwartz (1986) combine both the global and national factors in their empirical assessment of integration versus segmentation within a CAPM framework. They augment the international CAPM with a pure national systematic risk, and the domestic CAPM with a pure global systematic risk as follows:

$$E(R_i) = \gamma_0 + \gamma_1 \beta_i^G + \gamma_2 \beta_i^{N-G} \qquad (1)$$

and

$$E(R_i) = \delta_0 + \delta_1 \beta_i^N + \delta_2 \beta_i^{G-N} \qquad (1')$$

with implicit restrictions of $\gamma_1 \equiv E(R_G) - \gamma_0$ and $\delta_1 \equiv E(R_N) - \delta_0$ respectively. Here, the subscript "*i*" indicates a specific sectoral portfolio. For this portfolio, β_i^G and β_i^N are the systematic risks relative to the global and national indices, whose excess returns are represented by $R_G \equiv R_G^* - R_f$ and $R_N \equiv R_N^* - R_f$, with R_G^* and R_N^* being, respectively, the global and national nominal returns and R_f a risk-free rate. The terms, $\boldsymbol{\beta}_i^{N-G}$ and $\boldsymbol{\beta}_i^{G-N}$, represent systematic risks relative to hypothetical "pure" national and global indices. A "pure" national (/global) index is notionally one whose returns are independent of returns on the global (/national) index. Empirical counterparts to these hypothetical indices are constructed by a least squares decomposition: an OLS regression of the national (/global) market return upon the global (/national) returns produces residuals that are uncorrelated with the global (/national) returns and thus serve as "pure" national (/global) returns.

Eq. (1) reduces to the international CAPM when $\gamma_2 = 0$. A null hypothesis of integration can therefore be tested against the alternative of segmentation as $H_0 : \gamma_2 = 0$ vs. $H_A : \gamma_2 > 0$ in Eq. (1). Similarly a null of segmentation can be tested as the hypothesis that Eq. (1′) reduces to a domestic CAPM, viz: $H_0 : \delta_2 = 0$ vs. $H_A : \delta_2 > 0$. Jorion and Schwartz (1986) also note that Eqs. (1) and (1′) can be interpreted as a two-factor version of Ross's (1976) arbitrage pricing theory. Hence, as suggested by more recent literature such as Hardouvelis *et al.* (2006), stock markets may be partially integrated when both the global and national systematic risks are significantly and positively priced in Eqs. (1) and (1′).

In order to implement the test for integration, Jorion and Schwartz (1986) decompose the rate of return on portfolio *i* into an expected component and an innovation consisting of shocks to the global market, shocks to the pure national market and asset-specific disturbances as follows:

$$R_{it} = E(R_{it}) + \beta_i^G (R_{Gt} - E(R_{Gt})) + \beta_i^{N-G} \omega_{(N-G)t} + e_{it}^G \qquad (2)$$

where the expected value of each element of the innovation is zero by construction. Under rational expectations, Eq. (1) can be substituted into Eq. (2) to obtain Eq. (3) as follows.

$$R_{it} = \gamma_0 (1 - \beta_i^G) + \gamma_2 \beta_i^{N-G} + \beta_i^G R_{Gt} + \beta_i^{N-G} \omega_{(N-G)t} + e_{it}^G \qquad (3)$$

Under the joint hypothesis of integration and efficiency of the global market, γ_2 should be equal to zero. The test equation for the segmentation hypothesis, $H_0 : \delta_2 = 0$, can be devised similarly, from Eqs. (1′) and (2′):

$$R_{it} = \delta_0 (1 - \beta_i^N) + \delta_2 \beta_i^{G-N} + \beta_i^N R_{Nt} + \beta_i^{G-N} \omega_{(G-N)t} + e_{it}^N \qquad (3')$$

In the light of the considerable literature on time-varying systematic risk, we modify Eq. (3) as follows.

$$R_t = \mu_t + \beta_t^G R_{Gt} + \beta_t^{N-G}\omega_{(N-G)t} + e_t^G \quad (3.1)$$

$$\mu_t = \mu_{t-1} + \varepsilon_t \quad (3.1.1)$$

$$\beta_t^G = \beta_{t-1}^G + \eta_t \quad (3.1.2)$$

$$\beta_t^{N-G} = \beta_{t-1}^{N-G} + \upsilon_t \quad (3.1.3)$$

We allow the parameters, μ_i, β_i^G and β_i^{N-G}, of Eq. (3) to be time-varying by following independent random walks. The restriction, $\mu_t = \gamma_0(1-\beta_t^G) + \gamma_2\beta_t^{N-G}$, is now a testable implication of the asset pricing theory that underpins Eq. (3). Eqs. (3.1.1) to (3.1.3) constitute the transition equations of a state space model with Eq. (3.1) as the measurement equation. Similarly, a set of transition equations and a measurement equation can be associated with Eq. (3′). The Kalman smoothing technique can then be used to estimate series for the time-varying parameters in these two state space models.

There are two advantages to assuming that the parameters of the measurement equation follow a random walk. Firstly, it avoids model specification error when generating the time-varying coefficients with more explicit structural models. Some studies, such as Bekaert and Harvey (1995) and Hardouvelis *et al.* (2006), derive time-varying coefficients through instrumental variable methods rather than the state space model we use in this study. Harvey (2001) argues that the estimates of time-varying coefficients obtained using instrumental variables are very sensitive to the choice of instruments and are thus liable to suffer from omitted variable bias. Secondly, the random walk assumption removes the need to use leads and lags of the excess returns of the national and global indices in order to correct for thin trading effects, as in Jorion and Schwartz (1986) and Wang and Di Iorio (2007).

Once we have obtained the series of values for these time-varying parameters, we can test for integration and segmentation, respectively, in the following two-factor asset pricing models[2]:

$$\mu_t = \gamma_0(1-\beta_t^G) + \gamma_2\beta_t^{N-G} + \xi_t^G \quad (3.2)$$

and

$$\mu_t = \delta_0(1-\beta_t^N) + \delta_2\beta_t^{G-N} + \xi_t^N \quad (3.2')$$

A joint hypothesis of integration and efficiency of the global market implies $\gamma_2 = 0$ in Eq. (3.2), whilst, in Eq. (3.2′), $\delta_2 = 0$ under the joint hypothesis of segmentation and efficiency of the national market.

[2] Such a two-step approach is widely used in empirical studies but we note that it leads to some loss of efficiency since sampling error from the estimation in the first stage is disregarded in the second stage. When interpreting results, we should bear in mind that the reported standard errors may therefore be underestimated.

3.2. Impact of financial crises on asset pricing

In order to examine the impact of crises on stock market integration, we will apply two approaches to the estimation of Eqs. (3.2) and (3.2′). Firstly, we introduce an indicator for the recent global financial crisis: a dummy variable, D_t, taking a value of unity from August 2007 to March 2009 and set to zero outside of this period. This indicator variable is interacted with the regressors in Eqs. (3.2) and (3.2′) so as to permit a structural break, as follows:

$$\boldsymbol{\mu_t} = \gamma_0(1-\boldsymbol{\beta_t^G}) + \gamma_2\boldsymbol{\beta_t^{N-G}} + \boldsymbol{c}_{\gamma 1}\boldsymbol{D_t}(1-\boldsymbol{\beta_t^G}) + \boldsymbol{c}_{\gamma 2}\boldsymbol{D_t}\boldsymbol{\beta_t^{N-G}} + \boldsymbol{\xi_t^G} \qquad (3.3)$$

and

$$\boldsymbol{\mu_t} = \boldsymbol{\delta}_0(1-\boldsymbol{\beta_t^N}) + \boldsymbol{\delta}_2\boldsymbol{\beta_t^{G-N}} + \boldsymbol{c}_{\delta 1}\boldsymbol{D_t}(1-\boldsymbol{\beta_t^N}) + \boldsymbol{c}_{\delta 2}\boldsymbol{D_t}\boldsymbol{\beta_t^{G-N}} + \boldsymbol{\xi_t^N} \qquad (3.3')$$

Eqs. (3.3) and (3.3′) can be estimated by OLS. The impact of the recent financial crisis on the roles of the national and global systematic risks in China’s asset pricing can be captured by testing hypotheses relating to the coefficients on the interaction terms. A hypothesis of integration during the crisis period is represented by $\boldsymbol{\gamma}_2 + \boldsymbol{c}_{\gamma 2} = 0$ in Eq. (3.3), whilst segmentation in the crisis period can be stated by the hypothesis $\boldsymbol{\delta}_2 + \boldsymbol{c}_{\delta 2} = 0$ in Eq. (3.3′).

As an alternative approach, we use a two-state regime-switching technique to distinguish endogenously between periods of high volatility, representing crises, and those of low volatility that the Chinese sectors may have experienced during the sample period under study. This approach permits separate parameter values in the two discrete regimes, thus distinguishing market behaviour between turbulent and tranquil episodes. Following Hamilton (1989, 1990), we assume that, at each observation date, the parameters in Eqs. (3.2) and (3.2′) take values appropriate to one of the two discrete regimes as follows.

$$\mu_t = \gamma_{0,s_t}(1-\beta_t^G) + \gamma_{2,s_t}\beta_t^{N-G} + \xi_t^G \qquad (3.4)$$

and

$$\mu_t = \delta_{0,s_t}(1-\beta_t^N) + \delta_{2,s_t}\beta_t^{G-N} + \xi_t^N \qquad (3.4')$$

where $\operatorname{var}(\xi_t^G) = \sigma_{Gs_t}^2$ and $\operatorname{var}(\xi_t^N) = \sigma_{Ns_t}^2$ with $s_t \in [1,2]$. Parameter values may vary across regimes but are constant within regimes. Regime-switching may occur at any point in time. Although the regimes are never directly observed, probabilistic statements can be made about the relative likelihood of their occurrence. The latent regime indicator, S_t, is modelled as a first-order Markov process: at each point in time, there is a possibility of transition to the alternative regime, the likelihood of this regime-switch occurring depends only upon which regime was in force for the immediately preceding observation. The transition probabilities are detailed as follows.

$$\Pr(\boldsymbol{S_t} = 1 | \boldsymbol{S_{t-1}} = 1) = \boldsymbol{P}$$
$$\Pr(\boldsymbol{S_t} = 2 | \boldsymbol{S_{t-1}} = 1) = 1 - \boldsymbol{P}$$
$$\Pr(\boldsymbol{S_t} = 2 | \boldsymbol{S_{t-1}} = 2) = \boldsymbol{Q}$$
$$\Pr(\boldsymbol{S_t} = 1 | \boldsymbol{S_{t-1}} = 2) = 1 - \boldsymbol{Q}$$

How the Chinese portfolio returns evolve between these two states is fully described by the mean and variance parameters noted above and the two transition probabilities, P (probability of staying in state 1) and Q (probability of staying in state 2). All these parameters are estimated simultaneously using a maximum likelihood approach. The estimated variances can then be used to identify the two states: a high-volatility state of financial crisis and a low-volatility state of tranquil periods. The estimated parameters are used to calculate $\{p_t\}$, a series of smoothed estimated probabilities for regime 1 being in force at each point in time. Ang and Bekaert (2002) have devised a regime classification measure (RCM) as follows to measure the quality of the two-state regime classification.

$$\boldsymbol{RCM} = 400 \times \frac{1}{\boldsymbol{T}} \sum_{\boldsymbol{t}=1}^{\boldsymbol{T}} \boldsymbol{p_t}(1 - \boldsymbol{p_t}) \qquad (4)$$

A model that clearly identifies the more likely regime at each point in time will have a RCM close to zero whilst a model that cannot distinguish between regimes will produce a RCM close to 100.

4. Data and preliminary analysis

Data for this study are monthly time series. The sample period[3] is from January 1997 until March 2013, avoiding the extremely volatile period of early development of China's stock market. The global market portfolio is represented by the MSCI world all country index. The national market index is represented by the China A-share market index compiled by Datastream, as it covers the vast majority of stock listings on China's main stock exchanges in Shanghai and Shenzhen. We also downloaded ten A-share sector indices from Datastream as the Chinese portfolio indices in this study. The ten sectors are: Basic Materials, Consumer Goods, Consumer Services, Financials, Healthcare, Industrials, Oil and Gas, Technology, Telecom, and Utilities. All the Chinese indices are converted into US dollars, using dollar/yuan exchange rates obtained from the website of the Federal Reserve. We use the US 30-day Treasury bill return, obtained from Kenneth French's website[4], as the risk-free rate

[3] The sample periods for sectors of Oil & Gas and Technology are April 1998-March 2013 and April 1997-March 2013 respectively.

[4] http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html.

for the world index and the interest rate for the Chinese 3-month time deposit from Datastream as the risk-free rate for all the Chinese indices. The excess returns of the global index and the Chinese national and portfolio indices are calculated as the realised returns minus their respective risk-free rate.

Table 1 presents summary statistics of the returns under study. For the full sample period, the returns on the Chinese national ("A-share") market index and also on the ten Chinese sectoral indices have higher standard deviation than does the global ("MSCI World AC") market index. This relatively high volatility in the Chinese markets is compensated by higher average returns, except for three of the sectoral indices. During the crisis period, all of these equity indices exhibit increased variance. With the exception of the Telecom sectoral index, they all also experience a reduction in mean returns. The variance increase is higher for the Chinese markets than for the global index, compensated by smaller reductions in mean returns. The correlation between the Chinese A-share index and the MSCI world all country index in the crisis period is lower than that in whole period under study, contrary to the expectation of Longin and Solnik (1996) and the calculation by Ang and Chen (2002) that stock market correlations increase in times of financial crises.

Table 1 Summary statistics of the returns

	1997.1 – 2013.03		2007.8-2009.03	
	Mean	Standard deviation	Mean	Standard deviation
MSCI World AC index	0.287	5.262	-4.060	7.458
US risk-free rate	0.215	0.176	0.169	0.118
China A market index	0.322	8.589	-1.986	14.390
Chinese risk-free rate	2.233	0.595	2.844	0.619
Basic materials	0.207	10.664	-2.378	19.775
Consumer goods	0.384	9.429	-1.708	16.737
Consumer services	0.526	9.840	-2.390	17.293
Financials	0.416	10.022	-1.733	13.223
Healthcare	0.527	9.246	-0.327	14.671
Industrials	0.072	10.254	-2.004	16.393
Oil & Gas	0.483[a]	9.543[a]	-1.839	17.806
Technology	0.537[b]	10.802[b]	-1.389	18.897
Telecom	-0.189	11.918	-0.079	15.734
Utilities	0.334	9.183	-1.114	15.556
Correlation coefficient between World AC and China A market indices	0.17271		0.05409	

Note: [a] Sample period is April 1998-March 2013.
[b] Sample period is April 1997-March 2013.

5. Empirical results

5.1. Estimates of the global and national systematic risks

We first estimate by OLS, for each of the Chinese sector portfolios, constant parameter versions of Eq. (3.1) - an international CAPM augmented with the component of national excess return that is independent of the global market, and of Eq. (3.1′) - a domestic CAPM augmented with the component of global excess return that is independent of the national market. The OLS coefficients, reported in Table 2, serve as estimates for a set of benchmark time-invariant systematic risks: β^{G}, $\beta^{N\text{-}G}$, β^{N} and $\beta^{G\text{-}N}$. For all sectors, the estimated parameters of Eq. (3.1′) favour a null hypothesis of a simple domestic CAPM $(\boldsymbol{\beta}^{\boldsymbol{G-N}} = 0)$ rather than the augmented CAPM alternative $(\boldsymbol{\beta}^{\boldsymbol{G-N}} > 0)$. In the augmented international CAPM, Eq. (3.1), the purely national excess return is significant and positive for all sectors, rejecting a pure international CAPM. The estimated domestic effects, β^{N} and $\beta^{N\text{-}G}$, are similar in magnitude and are positively signed; they are larger than the estimated global effects, β^{G} and $\beta^{G\text{-}N}$. The balance of evidence obtained by this "constant-betas" approach suggests that domestic factors are essentially relevant to the pricing of Chinese equities.

To obtain time-varying systematic risks, we estimate Eqs. (3.1) and (3.1′) using the Kalman smoothing technique, and suppressing the first twenty data points in the estimated series, in order to avoid the instability and extreme values that are typical in the initialisation period of the smoothing algorithm. The time-varying systematic risks obtained from Eqs. (3.1) and (3.1′) are graphed respectively in Figures 1 and 2.

These figures show that the national systematic risk varies relatively more, especially in the sectors of Financials, Industrials, Telecom and Technology, while the global systematic risk is relatively stable over time. The national systematic risk is confirmed to be greater than the global systematic risk in all cases. The estimates of the time-varying national systematic risk, β^{N}, in Eq. (3.1′) are approximately the same as the purely national systematic risk term, $\beta^{N\text{-}G}$, in Eq. (3.1). Given that β^{N} includes the component of the global excess return that is not orthogonal to the national excess return, we interpret the same magnitude of β^{N} and $\beta^{N\text{-}G}$ as suggesting that such global excess return may have minimal influence on the excess returns in the Chinese markets. The estimates of the global systematic risk, β^{G}, of Eq. (3.1) are about 0.2 units higher than the estimated purely global systematic risk, $\beta^{G\text{-}N}$, of Eq. (3.1′), suggesting that the component part of the national excess return which is not orthogonal to the global excess return is amplifying the effect of the global excess return.

These conclusions reinforce the impression, gained earlier from OLS estimation of the constant-parameter models, that the national excess returns play a more significant role than the global one in the pricing of Chinese stocks. Since most of the Chinese industry-based indices are less sensitive to the changes in the global market return, there appears to be some suggestion in the time-varying betas that these sectors may provide international investors with an opportunity for risk diversification.

Table 2 Estimates of Eqs. (3.1) and (3.1′) by OLS during 1997.1-2013.3

	Eq. (3.1)			Eq.(3.1′)		
	μ	β^{G}	$\beta^{N\text{-}G}$	μ	β^{N}	$\beta^{G\text{-}N}$
Basic	-2.054***	0.383***	1.149***	0.182	1.155***	0.055
Materials	(0.285)	(0.054)	(0.033)	(0.292)	(0.033)	(0.055)
Consumer	-1.860***	0.158**	0.969***	-0.020	0.957***	-0.119*
Goods	(0.334)	(0.064)	(0.039)	(0.342)	(0.038)	(0.065)
Consumer	-1.731***	0.329***	1.002***	0.217	1.007***	0.043
Services	(0.339)	(0.065)	(0.040)	(0.347)	(0.039)	(0.066)
Financials	-1.837***	0.272***	0.995***	0.082	0.994***	-0.012
	(0.372)	(0.071)	(0.044)	(0.381)	(0.043)	(0.072)
Healthcare	-1.715***	0.128	0.803***	-0.190	0.793***	-0.101
	(0.453)	(0.086)	(0.053)	(0.464)	(0.052)	(0.088)
Industrials	-2.184***	0.324***	1.030***	-0.187	1.033***	0.030
	(0.360)	(0.069)	(0.042)	(0.369)	(0.042)	(0.070)
Oil&Gas[a]	-1.714***	0.342***	0.929***	0.101	0.937***	0.076
	(0.402)	(0.075)	(0.048)	(0.412)	(0.047)	(0.076)
Technology[b]	-1.694***	0.133	0.989***	0.176	0.973***	-0.155
	(0.522)	(0.099)	(0.063)	(0.535)	(0.062)	(0.100)
Telecom	-2.437***	0.211*	0.861***	-0.783	0.857***	-0.035
	(0.671)	(0.128)	(0.079)	(0.687)	(0.077)	(0.130)
Utilities	-1.907***	0.114*	0.926***	-0.158	0.911***	-0.150**
	(0.339)	(0.065)	(0.040)	(0.347)	(0.039)	(0.066)

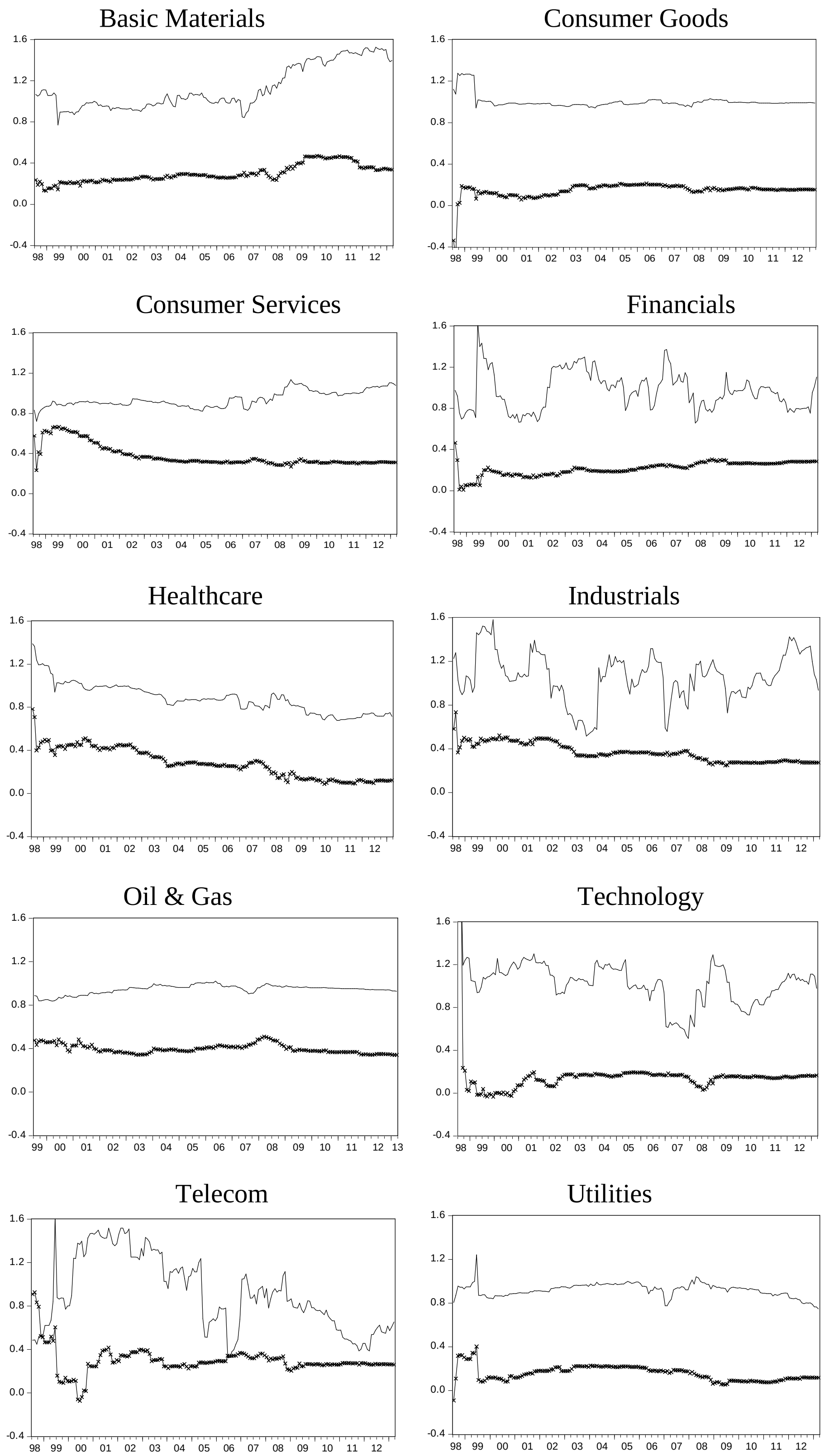
Note: This table reports estimates of Eqs.

(3.1): $\boldsymbol{R}_t = \boldsymbol{\mu} + \boldsymbol{\beta}^{G}\boldsymbol{R}_{Gt} + \boldsymbol{\beta}^{N-G}\boldsymbol{\omega}_{(N-G)t} + \boldsymbol{e}_t$ where $\underset{\text{s.e.}}{\boldsymbol{\omega}_{(N-G)t}} = \boldsymbol{R}_{Nt} + \underset{(0.615)}{1.932} - \underset{(0.117)}{0.286}\boldsymbol{R}_{Gt}$

(3.1’): $\boldsymbol{R}_t = \boldsymbol{\mu} + \boldsymbol{\beta}^{N}\boldsymbol{R}_{Nt} + \boldsymbol{\beta}^{G-N}\boldsymbol{\omega}_{(G-N)t} + \boldsymbol{e}_t$ where $\underset{\text{s. e.}}{\boldsymbol{\omega}_{(G-N)t}} = \boldsymbol{R}_{Gt} - \underset{(0.381)}{0.272} - \underset{(0.043)}{0.105}\boldsymbol{R}_{Nt}$

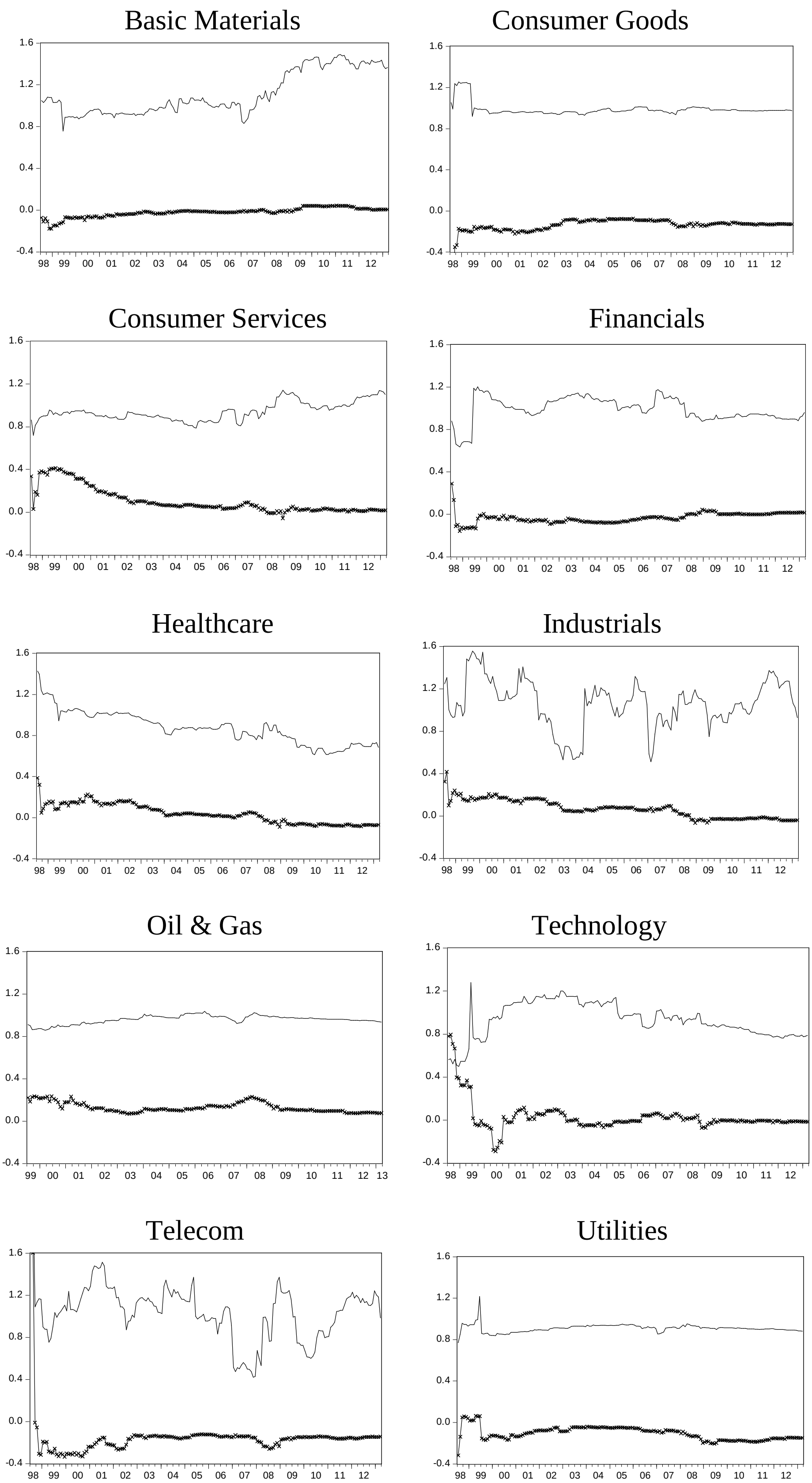
The heteroscedasticity-consistent (Eicker-White) standard errors are in brackets.
***, ** and * represent levels of significance at 1%, 5% and 10% respectively.
[a] Sample period is 1998.4-2013.3.
[b] Sample period is 1997.4-2013.3.

Figure 1 Time-varying global systematic risk and purely national systematic risk



Note: Lines in –x-x-x- represent the global systematic risk and those in — are the purely national systematic risk. These are estimated from Eq. (3.1) by the Kalman smoothing technique.

Figure 2 Time-varying national systematic risk and purely global systematic risk



Note: Lines in –x-x-x- represent the purely global systematic risk and those in — are the national systematic risk. These are estimated from Eq. (3.1') by the Kalman smoothing technique.

5.2. Tests for integration versus segmentation

In order to formally test for integration versus segmentation during the period under study, we substitute the time-varying coefficients of Eqs. (3.1) and (3.1′), respectively, into Eqs. (3.2) and (3.2′) and control for the recent financial crisis using a dummy variable as in Eqs. (3.3) and (3.3′). Because the state-space model restricts the time-varying coefficients to be random walks, there is a hypothetical risk of obtaining spurious regressions when estimating whether or not the CAPM-implied relations between these parameters are empirically supported. Unit root tests were therefore applied to the residuals from OLS estimates of Eqs. (3.2) and (3.2′). The testing framework was ADF without intercept or trend and with lag-lengths auto-selected by the software package. Of the 20 tests, only 2 accepted the null hypothesis of a unit root in the residuals at 10% significance, 2 accepted the null at 5% significance and the remainder rejected the null at the 5% significance level or better. This was taken to be evidence that fitted versions of Eqs. (3.2) and (3.2′) and therefore also of Eqs. (3.3) and (3.3′) could be treated as non-spurious regressions.

5.2.1. Test results in the period excluding the global financial crisis

The estimated results of Eqs. (3.3) and (3.3′) are reported in Table 3. These results confirm that an analysis with time-varying betas may lead to conclusions that differ from those obtained under a fixed-beta constraint. We have seen that the analysis with fixed betas favoured a simple domestic CAPM, implicitly rejecting a market integration hypothesis, and did so for all sectors. Now, with time-varying betas, we find that the results are less conclusive and show more variation between sectors.

With time-varying betas, the segmentation hypothesis requires $\delta_2 = 0$ in Eq. (3.3′), but we find this only for the Consumer Goods sector. In this sector, the null hypothesis of $\delta_2 = 0$ (insignificant exposure to the purely global systematic risk) is not rejected and the coefficient, δ_0, of the national systematic risk is correctly signed[5] and statistically significant, suggesting that the national systematic risk is the only significant factor in its asset pricing. The Consumer Goods sector is hence segmented with the global market. In the other sectors there is evidence of international influence on China's asset pricing, albeit with an unexpected negative sign in some cases. For example, we reject the null hypothesis of $\delta_2 = 0$ in favour of positive exposure to the purely global systematic risk ($\delta_2 > 0$) in the sectors of Basic Materials, Financials, Healthcare, Oil and Gas and Utilities. The coefficient of the national systematic risk, δ_0, is further correctly signed and statistically significant in all these sectors except Utilities. Hence

[5] The correct sign for the coefficients of $(1-\beta_t^G)$ and $(1-\beta_t^N)$ in Eqs. (3.3) and (3.3'), γ_0 and δ_0, must be negative in order for the global and national systematic risks to be positively priced respectively.

sectors such as Basic Materials, Financials, Healthcare and Oil and Gas are 'partially integrated' with the global market as in Hardouvelis *et al.* (2006).

At the same time, the complete integration hypothesis, i.e. $\gamma_2 = 0$ in Eq. (3.3), is supported in only two sectors: Healthcare and Utilities. In these two sectors, the null hypothesis of $\gamma_2 = 0$ (insignificant exposure to the purely domestic systematic risk) is not rejected and the coefficient, γ_0, of the global systematic risk is correctly signed and statistically significant, suggesting that the global systematic risk is the only significant factor in asset pricing of these sectors. In the other sectors, purely domestic influences are relevant to pricing, albeit again with a negative sign in some cases. Of these sectors, Financials and Oil and Gas are confirmed to be partially integrated with the global market as they are exposed positively to both systematic risks in Eq. (3.3).

Overall, Table 3 suggests that the polarised extremes, "integration" or "segmentation", are evident in only a small number of sectors. In the majority of cases, both the purely global and domestic factors seemingly contribute to asset pricing, pointing to a possible conclusion of partial integration and thus leaving some opportunity for useful diversification of international portfolios by inclusion of Chinese equity assets - though this may require a considered selection of sectors in which to invest.

The preceding conclusion ignores the question mark raised by the instances of apparently negative risk pricing, e.g., negative pricing of both the purely national and global systematic risks in the sectors of Consumer Services, Industrials, and Technology. Negatively priced risks suggest that investors are unwilling to pay any positive premiums to hedge against global downturns and/or national downturns in these sectors. Although this result is inconsistent with the theory of risk aversion, it is not uncommon to observe negatively priced systematic risks, especially the national systematic one, in either developed or emerging markets. Mo and Wu (2007), in their study of the US, UK and Japan, find that all the three markets charge a positive price for the global return risk, but the country-specific risk is negatively priced in the case of Japan. Tai (2007) and Omran (2007) also find a negatively priced local market risk in the case of Egypt and in the cases of India and Malaysia respectively. Haugan (1999) attributes the negative trade-off between return and beta to the omission of a size or liquidity premium effect in the CAPM model while De Santis and Gerard (1997) show that negative price of market risk exists during those years when interest rates and inflation are unusually high. In the case of China, there may be further institutional factors that need to be acknowledged in models of asset pricing.

Table 3 Estimates of Eqs. (3.3) and (3.3') during 1998.7-2013.3 by OLS

	Basic Materials	Consumer Goods	Consumer Services	Financials	Healthcare	Industrials	Oil&Gas[a]	Technology	Telecom	Utilities
Eq. (3.3): Test of integration against segmentation										
γ_0	-2.364*** (0.059)	-1.492*** (0.463)	3.142*** (0.524)	-4.001*** (0.121)	-2.113*** (0.148)	-2.719*** (0.375)	-6.499*** (0.849)	-0.578*** (0.150)	-3.603*** (0.219)	-1.844*** (0.175)
γ_2	-0.211*** (0.037)	-0.688** (0.395)	-4.446*** (0.349)	0.497*** (0.096)	-0.126 (0.119)	-0.733*** (0.217)	2.461*** (0.551)	-0.411*** (0.123)	0.270** (0.160)	-0.076 (0.161)
$c_{\gamma 0}$	0.976*** (0.390)	3.167 (4.316)	-3.271 (3.365)	0.051 (0.681)	3.035* (2.225)	-3.714* (3.279)	6.567*** (2.700)	-0.853* (0.648)	2.046 (1.679)	1.608* (1.152)
$c_{\gamma 2}$	-0.456** (0.229)	-2.760 (3.659)	2.845 (2.323)	0.037 (0.508)	-2.980* (2.077)	3.089* (2.120)	-3.901** (1.534)	0.227 (0.624)	-1.840* (1.287)	-1.627* (1.032)
$\gamma_0 + c_{\gamma 0}$	-1.388** (0.386)	1.675 (4.291)	-0.129 (3.324)	-3.950*** (0.607)	0.922 (2.220)	-6.433** (3.257)	0.068 (2.562)	-1.431** (0.637)	-1.558 (1.665)	-0.237 (1.138)
$\gamma_2 + c_{\gamma 2}$	-0.667*** (0.226)	-3.448 (3.673)	-1.601 (2.296)	0.534 (0.498)	-3.107* (2.073)	2.356 (2.109)	-1.439 (1.431)	-0.133 (0.611)	-1.570 (1.277)	-1.703** (1.020)
Eq. (3.3'): Test of segmentation against integration										
δ_0	-1.389*** (0.061)	-4.297*** (0.551)	-1.567*** (0.442)	-3.319*** (0.337)	-0.599** (0.243)	-1.232*** (0.255)	-8.599*** (0.694)	-2.896*** (0.318)	-3.164*** (0.254)	3.433*** (0.545)
δ_2	1.785*** (0.284)	0.078 (0.222)	-3.495*** (0.278)	7.314*** (0.644)	3.891*** (0.451)	-4.909*** (0.545)	4.441*** (0.381)	-2.755*** (0.317)	-0.988*** (0.306)	2.461*** (0.436)
$c_{\delta 0}$	0.562** (0.254)	3.352 (4.081)	1.350 (1.595)	-0.166 (2.383)	-0.521 (0.729)	-0.437 (1.787)	9.327*** (3.619)	3.060*** (0.784)	-4.172*** (1.606)	-4.346* (2.722)
$c_{\delta 2}$	-12.47*** (3.124)	0.708 (1.015)	9.862*** (3.527)	2.768 (6.440)	-4.078* (2.857)	4.463 (4.108)	-1.527** (0.662)	-0.227 (1.243)	-5.514** (2.918)	-2.508* (1.600)
$\delta_0 + c_{\delta 0}$	-0.827*** (0.247)	-0.944 (4.044)	-0.216 (1.532)	-3.485* (2.359)	-1.120* (0.690)	-1.668 (1.768)	0.728 (3.552)	0.164 (0.716)	-7.335*** (1.586)	-0.913 (2.671)
$\delta_2 + c_{\delta 2}$	-10.68*** (3.111)	0.786 (0.990)	6.366** (3.516)	10.10* (6.407)	-0.187 (2.822)	-0.447 (4.072)	2.914*** (0.541)	-2.982*** (1.201)	-6.502** (2.902)	-0.047 (1.539)

Note: [a] Sample period is 1999.1-2013.3.
Standard errors are in brackets. ***, ** and * represent levels of significance at 1%, 5% and 10% respectively for a one-tailed test.

5.2.2. Impact on asset pricing of the recent global financial crisis

Table 3 also reports that estimates of the coefficients of the interaction terms, $c_{\gamma 0}$, $c_{\delta 0}$, $\boldsymbol{c}_{\gamma 2}$ and $\boldsymbol{c}_{\delta 2}$, are statistically significant in more than half of the sectors under study, suggesting changes in the sectoral exposures to the national and global systematic risks between the period of the global financial crisis (August 2007 - March 2009) and the period outside the crisis. Among the sectors with the significant differential coefficients, the global financial crisis tends to reduce more than it does to increase the sectoral exposures to the national and global systematic risks. Consequently, six of the ten sectors under study are no longer exposed to the global or national systematic risks ($\gamma_0 + c_{\gamma 0} = 0$ or $\delta_0 + c_{\delta 0} = 0$) in the crisis although the remaining four sectors continue to positively price the systematic risks respectively. Similarly, the null hypothesis of $\gamma_2 + c_{\gamma 2} = 0$ (insignificant exposure to the purely national systematic risk) cannot be rejected in seven of the ten sectors during the global financial crisis in Eq. (3.3) while the null hypothesis of $\delta_2 + c_{\delta 2} = 0$ (insignificant exposure to the purely global systematic risk) are not rejected in four of the ten sectors in Eq. (3.3′).

However, decreases in exposures to the systematic risks result in only one case where asset pricing is related to neither the national nor global systematic risk in the crisis. That is, the asset pricing of Consumer Goods mainly depends on risk specific to the sector, i.e., its idiosyncratic risk, in the crisis. In contrast with the three instances of apparently negative pricing of both the national and global systematic risks in the period excluding the financial crisis, there is only one case, namely the sector of Basic Materials, of such incidence in the crisis. Furthermore, the Chinese sectors, especially those prone to cyclical fluctuations, seem to be exposed relatively more to the purely global systematic risk than to the national one, emphasizing the extreme case of complete integration in the crisis. According to the results of Eq. (3.3) in Table 3, in the sectors such as Financials, Industrials and Technology, the null hypothesis of integration ($\gamma_2 + c_{\gamma 2} = 0$) cannot be rejected while the coefficient of the global systematic risk ($\gamma_0 + c_{\gamma 0}$) is correctly signed and statistically significant. These three sectors become fully integrated with the global market during the crisis.

Meanwhile, there is no more case of partial integration in the crisis according to the results of Eq. (3.3) in Table 3. On the other hand, although the sector of Financials remains to be partially integrated with the global market in the crisis on the basis of Eq. (3.3′) in Table 3, two more sectors, namely Consumer Services and Oil and Gas, are found to be integrated with the global market in the crisis. In these two sectors, the null hypothesis of segmentation $(\delta_2 + c_{\delta 2} = 0)$ is rejected in favour of positive exposure to the pure global systematic risk $(\delta_2 + c_{\delta 2} > 0)$ and the coefficient of the national systematic risk ($\boldsymbol{\delta}_0 + \boldsymbol{c}_{\delta 0}$) is statistically insignificant.

Overall, we find stronger evidence of complete integration and weaker that of partial integration during the global financial crisis. The number of sectors that are integrated with the global market has increased to five in the global financial crisis from two in the period excluding the crisis, while the number of sectors that are partially integrated with the global market has decreased to one from five sectors. Our findings do seem to support the literature, such as Beine *et al.* (2010), that stock market integration strengthens in the global financial crisis.

Table 4 The Bai-Perron (2003) multiple change point analysis on Eqs. (3.2) and (3.2')

Sector	Eq. (3.2)		Eq. (3.2')	
	No. Breaks*	Time points	No. Breaks*	Time points
Basic Materials	3	2000:11 2008:01 2011:03	5	1999:09 2000:12 2007:01 2008:11 2011:05
Consumer Goods	5	1999:09 2001:02 2004:04 2006:04 2009:11	6	1999:09 2001:02 2003:12 2005:05 2007:01 2009:11
Consumer Services	5	1999:09 2002:06 2004:10 2007:01 2009:07	3	1999:12 2006:05 2010:01
Financials	1	2001:03	4	1999:09 2001:03 2002:06 2008:03
Healthcare	3	2000:07 2006:05 2008:02	4	1999:09 2000:12 2007:01 2009:02
Industrials	7	2000:05 2002:12 2004:04 2005:09 2007:01 2008:09 2010:08	7	2000:06 2003:05 2004:11 2007:01 2008:04 2010:09 2011:12
Oil&Gas	1	2000:06	3	2000:05 2003:01 2007:02
Technology	5	2000:02 2001:08 2002:11 2005:03 2007:12	2	2000:07 2006:04
Telecom	3	2000:02 2002:06 2008:01	6	1999:09 2000:12 2003:01 2005:05 2006:08 2008:09
Utilities	4	2000:06 2002:07 2005:09 2008:10	2	2000:06 2006:02

Note: * identified by using BIC.

5.3. *Stock market integration under the two-state regime*

As noted previously in Table 1, volatility of the returns on Chinese sectoral indices increased substantially during the recent global financial crisis. Within the full sample period, there could be other episodes of volatility crises. Using the Bai-Perron algorithm (Bai and Perron, 2003), we detect a range of 1 to 7 structural breaks, suggesting 2 to 8 regimes over time, in the relationships of Eqs. (3.2) and (3.2′) for the ten sectors under study. Table 4 further reports that the breaks have occurred at different time points within sectors.

Because the location and number of the change points are not uniform within sectors, we do not attempt to test for integration versus segmentation with multiple change points. Instead, we apply the Markov two-state regime-switching approach to the estimation of Eqs. (3.2) and (3.2′) as in Eqs. (3.4) and (3.4′). The Markov regime-switching technique characterises the potential for structural breaks as a consequence of occasional transition between two regimes, which we can index as a high-volatility state and a low-volatility state. This allows a hypothetical regularity in stock market behaviour to have different parameter values during turbulent and tranquil periods. The results of estimating Eqs. (3.4) and (3.4′) by the Markov two-state regime-switching technique for the period of July 1998-March 2013 are reported in Tables 5 and 6 respectively.

As shown in Tables 5 and 6, the regimes can be distinguished by different values for the standard deviation of the regression residuals, σ_{s_t}, which is a volatility measure. Regimes 1 and 2 are, respectively, the low-volatility and high-volatility states. There is noticeable persistence in both states for all the sectors since the probabilities of staying in the existing states, P(1,1) and 1-P(1,2), are both large. The finding of highly persistent regimes is not unusual; Schaller and van Norden (1997), Assoe (1998) and Liu *et al.* (2012) also report that both the low-variance and high-variance states are highly persistent in their studies on stock market returns. Furthermore, the regime classification measure (RCM) takes values close to its lower limit, suggesting that our estimated two-state regime-switching model is able to distinctly classify the two alternative modes of behaviour for Chinese portfolio returns.

Table 5 Estimates of Eq. (3.4) during 1998.7-2013.3 by the Markov two-state regime-switching technique

	Basic Materials	Consumer Goods	Consumer Services	Financials	Healthcare	Industrials	Oil&Gas[a]	Technology	Telecom	Utilities
γ_{01}	-0.126*** (0.025)	-1.501*** (0.474)	2.887*** (0.159)	-3.485*** (0.065)	-2.624*** (0.114)	-4.533*** (0.181)	-4.606*** (0.236)	-1.533*** (0.052)	-3.988*** (0.098)	-2.123*** (0.032)
γ_{21}	-0.047*** (0.015)	-1.058*** (0.157)	2.830** (1.274)	-4.187*** (0.381)	-2.390*** (0.086)	-1.345*** (0.244)	-8.774*** (1.263)	-0.560*** (0.183)	-1.475* (1.009)	-0.529*** (0.170)
γ_{02}	-2.316*** (0.038)	-0.884* (0.550)	-4.631*** (0.106)	0.159*** (0.051)	-0.188** (0.084)	-0.447*** (0.104)	1.392*** (0.153)	-0.044 (0.045)	0.369*** (0.074)	-0.022 (0.034)
γ_{22}	-3.069*** (0.030)	-0.302* (0.185)	-3.841*** (0.858)	0.426 (0.391)	0.593*** (0.077)	-0.881*** (0.144)	3.069*** (0.772)	-0.226 (0.199)	-0.687 (0.726)	-1.034*** (0.153)
σ_1	0.002*** (0.001)	0.031*** (0.007)	0.028*** (0.004)	0.008*** (0.001)	0.018*** (0.003)	0.061*** (0.010)	0.006*** (0.001)	0.009*** (0.001)	0.101*** (0.012)	0.010*** (0.001)
σ_2	0.006*** (0.001)	0.093*** (0.012)	0.712*** (0.109)	0.240*** (0.066)	0.057*** (0.010)	0.144*** (0.021)	0.403*** (0.094)	8.098*** (1.335)	1.960*** (0.512)	0.034*** (0.006)
P(1,1)	0.94*** (0.034)	0.989*** (0.011)	0.984*** (0.014)	0.993*** (0.011)	0.985*** (0.011)	0.984*** (0.013)	0.980*** (0.012)	0.995*** (0.005)	0.995*** (0.006)	0.994*** (0.007)
P(1,2)	0.02** (0.011)	0.012* (0.009)	0.008 (0.008)	0.019 (0.040)	0.007 (0.008)	0.008 (0.009)	0.064 (0.044)	0.007 (0.007)	0.011 (0.015)	0.006 (0.007)
RCM	0.742	1.278	1.795	3.841	0.914	1.252	2.569	1.037	1.732	4.325
No. iterations to achieve convergence	33	30	21	28	31	19	25	28	22	36
Function value	199.1424	-23.6127	-92.4100	109.3672	46.5663	-54.5225	101.6402	-47.7892	-101.8107	100.9375

Note: [a] Sample period is 1999.1-2013.3. ***, ** and * represent levels of significance at 1%, 5% and 10% respectively for a one-tailed test.
σ_1 and σ_2 are standard deviations of the regression residuals in regimes 1 and 2 respectively.
P(i,j) are estimated transition probabilities, for switching from regime j to regime i.
RCM is the regime classification measure, devised by Ang and Bekaert (2002) for the case of two states. A value of 0 means perfect classification while a value of 100 implies failure to classify.

Table 6 Estimates of Eq. (3.4') during 1998.7-2013.3 by the Markov two-state regime-switching technique

	Basic materials	Consumer goods	Consumer services	Financials	Healthcare	Industrials	Oil&Gas[a]	Technology	Telecom	Utilities
δ_{01}	-1.298*** (0.043)	-4.301*** (0.227)	-0.383*** (0.137)	-2.137*** (0.196)	0.254*** (0.045)	-0.838*** (0.098)	-5.147*** (0.250)	-0.518*** (0.075)	-3.622*** (0.107)	5.915*** (0.276)
δ_{21}	-1.744*** (0.104)	-4.407*** (0.992)	-3.518*** (1.245)	-5.646*** (1.001)	-3.750*** (0.262)	-1.398*** (0.426)	-7.979*** (1.236)	-1.487 (0.924)	-2.813*** (0.465)	-4.719*** (0.678)
δ_{02}	-1.058*** (0.295)	0.755*** (0.117)	-4.006*** (0.110)	8.730*** (0.353)	1.656*** (0.112)	-4.156*** (0.165)	4.770*** (0.094)	-6.159*** (0.109)	-0.833* (0.566)	3.458*** (0.248)
δ_{22}	2.418*** (0.423)	-0.186 (0.312)	-2.901*** (0.586)	2.957* (1.917)	5.616*** (0.384)	-5.281*** (0.971)	1.789** (0.775)	1.075* (0.564)	-1.281*** (0.452)	-1.913*** (0.542)
σ_1	0.004*** (0.001)	0.009*** (0.002)	0.016*** (0.002)	0.024*** (0.003)	0.006*** (0.001)	0.018** (0.004)	0.006*** (0.001)	0.060*** (0.007)	0.014*** (0.002)	0.015*** (0.002)
σ_2	0.043*** (0.007)	0.283*** (0.038)	0.565*** (0.098)	0.903*** (0.218)	0.132*** (0.019)	0.963*** (0.149)	0.429*** (0.097)	16.123*** (3.431)	0.443*** (0.061)	0.034*** (0.007)
P(1,1)	0.975*** (0.015)	0.917*** (0.038)	0.975*** (0.015)	0.988*** (0.009)	0.982*** (0.014)	0.947*** (0.029)	0.981*** (0.012)	0.978*** (0.012)	0.975*** (0.020)	0.977*** (0.012)
P(1,2)	0.023 (0.017)	0.033** (0.018)	0.036* (0.022)	0.038* (0.034)	0.017 (0.012)	0.047** (0.026)	0.062 (0.044)	0.074* (0.045)	0.017* (0.013)	0.046** (0.027)
RCM	7.598	10.279	5.191	3.931	2.057	11.424	2.954	3.656	5.015	10.443
No. iterations to achieve convergence	14	21	14	30	24	15	19	40	14	25
Function value	128.7700	-68.2732	-26.0441	11.8403	44.8440	-105.3906	98.7542	-118.9010	-75.8461	81.3459

Note: [a] Sample period is 1999.1-2013.3. ***, ** and * represent levels of significance at 1%, 5% and 10% respectively for a one-tailed test.
σ_1 and σ_2 are standard deviations of the regression residuals in regimes 1 and 2 respectively.
P(i,j) are estimated transition probabilities, for switching from regime j to regime i.
RCM is the regime classification measure, devised by Ang and Bekaert (2002) for the case of two states. A value of 0 means perfect classification while a value of 100 implies failure to classify.

When interpreting the test results, we still look for evidence of international influences on asset pricing at the Chinese sector level. We are interested in $H_0 : \gamma_{2s_t} = 0$ in Table 5, which supports the hypothesis of complete market integration, and $H_A : \delta_{2s_t} > 0$ in Table 6, which supports positive exposure to the purely global systematic risk. We are also interested in whether conclusions differ between low-volatility periods and high-volatility periods, expecting that these latter may have often been associated with some form of crisis.

According to the results of Eq. (3.4) in Table 5, there is no evidence of complete or partial integration during the low-volatility periods. We reject the null hypothesis of integration $(\gamma_{21} = 0)$ in favour of negatively-priced purely national systematic risk $(\gamma_{21} < 0)$ in all sectors but Consumer Services. In the Consumer Services sector, although the null hypothesis of $\gamma_{21} = 0$ is rejected in favour of $\gamma_{21} > 0$, the coefficient of the global systematic risk, γ_{01}, is wrongly signed, failing to support the conclusion that the Consumer Services sector is partially integrated with the global market. Similarly, we can reject the null hypothesis of segmentation $(\delta_{21} = 0)$ in favour of negatively priced purely global systematic risk $(\delta_{21} < 0)$ in all sectors but one according to the results of Eq. (3.4′) in Table 6. The exception here is the Technology sector, where the null hypothesis of segmentation is not rejected and the coefficient of the national systematic risk is correctly signed and statistically significant, leading to its segmentation from the global market in the low-volatility periods. Overall a striking observation is that the purely national and global systematic risks are, respectively, negatively priced in nine of the ten sectors under study during the low-volatility periods. It seems that investors do not wish to pay any positive premium to hedge against global and/or national downturns in the low-volatility periods.

In the high-volatility periods, there is no evidence of complete integration but rather evidence of partial integration. Although we do not reject the null hypothesis of integration $(\gamma_{22} = 0)$ in three sectors, namely Financials, Technology and Telecom, according to the results in Table 5, the coefficients of the global systematic risk, γ_{02}, in these three sectors are either incorrectly signed or statistically insignificant, failing to suggest a pure international CAPM for these cases. Meanwhile, only one sector, i.e., Healthcare, is found to be partially integrated with the global market as the null hypothesis of $\gamma_{22} = 0$ is rejected in favour of $\gamma_{22} > 0$ and its coefficient of the global systematic risk is correctly signed and statistically significant. On the other hand, there is more evidence of international influence on China's asset pricing in the high-volatility periods according to the results in Table 6. In these periods, we can reject the null hypothesis of segmentation $(\delta_{22} = 0)$ in favour of positive exposure to the purely global systematic risk $(\delta_{22} > 0)$ in five sectors. Of these, the sectors of Basic Materials and Technology are partially integrated with the global market, given their positive exposure to national systematic risk additionally. Overall, international influence on China's asset pricing becomes stronger, leading to partial integration in three sectors in the high-volatility periods as opposed to none in the low-volatility periods.

6. Conclusions

This paper investigates China's asset pricing mechanism with the objectives of assessing integration between the Chinese and global stock markets and detecting the impact of financial crises on such integration. The analysis with fixed betas supports a simple domestic CAPM, implicitly rejecting a market integration hypothesis, and does so for all sectors. With time-varying betas, we find that the results show more variation between sectors. While the polarized extremes of integration or segmentation are evident in a small number of sectors in the periods excluding the recent financial crisis, five of the ten sectors under study tend to expose to both global and domestic factors, possibly pointing to a conclusion of "partial integration" as in Hardouvelis *et al.* (2006) and thus leaving some opportunity for useful diversification of international portfolios by inclusion of Chinese equity assets. During the recent global financial crisis, the global systematic risk is the only significant factor in asset pricing of five sectors, leading to complete market integration in five of the ten sectors under study. These findings seem to support the notion that stock market integration strengthens in times of financial crises.

However, complete market integration appears to be a truly defining feature of the recent global financial crisis as the Chinese stock market is not found to be integrated with the global market in the high-volatility crisis periods using the two-state regime-switching approach. Nonetheless, the global systematic risk plays a greater role relative to the national systematic risk in asset pricing of the Chinese sectors in the high-volatility periods, leading to partial integration in three sectors in the high-volatility periods, as opposed to none in the low-volatility periods.

Finally, it is striking that the majority of the sectors under study have tended to negatively price either the purely global or national systematic risks in the low-volatility periods. Given that these results are inconsistent with the theory of risk aversion, the study of China's asset pricing could be extended to explore whether institutional features particular to China, such as sectoral variations in state ownership, can be usefully incorporated into this well-developed CAPM framework.

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