The regime-switching impact of global and regional financial markets on China's stock market

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Abstract

In this paper we examine to what extent the Chinese A-share market, allegedly fully protected from foreign influences by strict capital controls, is actually influenced by movements in international (New York) and regional (Hong Kong) markets. We study the longest available sample (April 1993 through October 2004) of active trading, with a weekly frequency for dollar indices. Before the East Asian crisis, the Shanghai A-share market index was cointegrated with the S&P500. This relationship broke down and a long run relationship with the Hang Seng index arose after the East Asian crisis. Second, we examine dynamic relationships within the framework of a Markov-switching vector error correction model. The nature of the error correction term changed over time. Prior to the East Asian crisis deviations form the long run relationship with New York led to sizeable error correction and lagged US returns had an impact on Chinese returns, while subsequently only deviations form the new long run relationship with Hong Kong mattered, with a lower speed of adjustment. The prospects of *de jure* financial opening and the growing awareness of valuation concepts (P/E ratio comparisons) among Chinese domestic investors help explain the recent evidence of financial integration in spite of capital controls.

JEL Classification Number F36, G15

Keywords: China's A-share market, Markov-switching VECM, international financial integration

I. Introduction

In the eyes of many observers recent changes in commodity markets and the real sector have dwarfed in importance the integration of China's financial sector into global financial markets (Drysdale and Huang, 2003) in as much as capital controls have been retained, preventing foreigners to access A-share markets and residents to access foreign stock markets. However since the end of 1999 internationalization has become one of the big themes of Chinese stock market reform. China made commitments at the time of its accession to the World Trade Organisation (WTO) but the latter, which focus mostly on the domestic banking sector, do not concern capital account liberalisation. However the opening of domestic capital markets went beyond WTO commitments. Thus foreign investors were allowed to acquire a share of domestic companies' capital and since late 2002 large foreign institutional investors have been allowed to invest in bonds or stocks listed in Chinese domestic markets under the Qualified Foreign Institutional Investors' (QFII) scheme. All these would represent potential channels for connexions with international markets.

On top of regulatory changes, over the last two years, the large current account imbalances at the global level have generated pressures towards RMB appreciation with associated "capital inflows" into China. There is thus a presumption that the integration of China's capital markets would have increased over the recent period. However casual observation implies that since the start the impact of international stock markets seems to have been important on China's stock market, possibly initially only through information channels and later on increasingly through hidden capital flows. Existing theoretical work and evidence on emerging markets show that external *de jure* financial openness is neither sufficient nor necessary for foreign stock prices to influence domestic stock prices (Aizenman, 2003). We will argue that for the international convergence of stock markets the thinking of investors may be more important than actual capital flows. When foreign stock markets become the norm in the mind of investors, for instance through levels of Price/Earnings ratios, the prospect of a future opening of domestic capital markets can become a strong vector of convergence of such markets towards international ones.

A puzzle to millions of individual Chinese domestic investors and to many observers is the everlasting bear phase of the secondary market since 2001. We want to assess to what extent the current state of international integration of China's stock market and its future

prospects can help explain away such a puzzle. More generally, we will examine how international stock markets impact on China's A-share market. We focus on New York as the major global stock market with which the A-share market has connections, and on Hong Kong as the regional market which is an important benchmark for Chinese domestic investors, given that up to two fifths of the Hang Seng index lists mainland Chinese firms. We consider to what extent the impact of these markets on the Shanghai market is different in the short and in the long run, and whether the anchor for the Chinese market has switched between the international and the regional markets depending on the period examined.

A consensus view among empirical researchers is that the Chinese stock market has been mostly protected from the influence of movements in foreign stock markets, both in the long run and in the short run. Existing work, using widely different samples (always ending in 2001 at the latest), does not find any evidence either of cointegration of A-share indices with foreign stock market indices, particularly New York and Hong Kong, or of dynamic effects of foreign returns on A-share returns. Neither does existing literature document any significant break in such relationships at the time of the East Asian crisis.

This paper first re-examines the issue of the existence of long run relationships with weekly data over the longest available sample of active trading in the A-share market (April 1993-October 2004). We pay particular attention to features neglected in previous work on China's stock market: the inclusion of a deterministic trend and the treatment of outliers, which are extremely large in the Chinese case. We provide fresh evidence for cointegration of the Shanghai A-share market index with either the Standard and Poor's 500 or the Hang Seng index respectively before and after the East Asian crisis. We consider the A-share index in dollar (as Kasa, 1992, as well as Phylaktis and Ravazollo, 2003) in order to account for the point of view of an international portfolio diversification strategy.

Second, we study dynamic relationships using a Markov-switching Vector Error Correction model. This leads us to consider to what extent the significance of the error correction term is regime-dependent. The specification used leaves room for a change in the anchor for the domestic index between international and regional indices.

In the next section we present existing empirical literature which has measured the degree of integration of emerging stock markets in general and the Chinese market in particular, with regional and global markets. In section three the methodology for Markov-switching vector error correction models will be introduced. Section four will present tests for the presence of long run relationships between the A-share market and global or regional

stock markets. In section five the results of the estimation of a Markov-switching vector error correction model will be provided. In section six we explain the background reasons for the evolving international integration of China's stock market. Section seven concludes.

II Assessing the degree of international integration of emerging equity markets

Existing work on emerging markets has provided evidence to the effect that external *de jure* financial openness is neither a sufficient nor a necessary condition for foreign stock prices to influence domestic stock prices. Thus, for East Asian emerging market countries, excluding China, Phylaktis and Ramazzolo (2004) have used multiple cointegrating techniques on Monthly data to examine, over the eighties and nineties, the time variation in the long run relationships between indices in emerging and developed stock markets before and after external financial liberalization. They find that the relaxation of ownership restrictions in the 1980s were not enough to integrate the Malaysian or Hong Kong markets with world markets, while, over the same period, the lack of removal of such restrictions in Taiwan and Thailand did not prevent the international integration of these two markets. The use of recursive cointegration tests leads them to conclude that the East Asian financial crisis did not change significantly the financial linkages of these emerging markets with world stock markets.

Recently many empirical studies have examined the case of China. Huang et al (2000) consider the causal and long run relationships between China and Hong Kong, Japan, Taiwan or the U.S. With daily data over October 1992 through June 1997, they do not find any evidence of a bivariate cointegrating relationship linking a Chinese market with either the US or any of the other three markets. They do not uncover any causal relationship from the US market (or any of the other three markets) to the Chinese market. Hsiao et al. (2003) using a linear VAR on first differences, with daily data over a recent, but short, period (September 2001 to December 2002), do not find either any causal relationship from the US market to the Chinese market, and with impulse response functions, show that a fall in US stock prices has no impact on Chinese stock prices.

Hatemi and Roca (2004) consider the interdependence between China, Hong Kong, Singapore and Taiwan with daily MSCI indices (which control for the absence of double listing) expressed in dollars, over the period January 1993-September 2001. Using long term causality test, they also are unable to document that the Hong Kong stock market is influential

on mainland China. They uncover very weak evidence (at the ten per cent level) that the US stock market started influencing China after the East Asian crisis. Goenewold et al. (2003) consider the relationships between mainland Chinese markets and the Hong Kong and Taiwanese markets with daily data over the October 1992 through November 2001 period. They confirm that Hong Kong is not cointegrated with China's A share markets, and that, in a VAR in first differences, such A-markets are isolated from the Hong Kong or Taiwanese markets.

Bahng and Shin (2003) focus on relationships between A-share markets and North East Asian markets (Japan and Korea), as well as the NYSE composite index. Considering daily data over the period from January 1991 through December 2000, they are also unable to provide evidence for a cointegrating vector between a Chinese market and the three foreign markets. In a VAR in first differences, with impulse response functions, they find no effect of US shocks on Chinese returns. Wang and Firth (2004), using daily data over the period November 1994 through September 2001, similarly conclude that mainland Chinese markets are affected neither by contemporaneous not by delayed bad news from the US market. They provide some evidence in favour of a structural change at the time of the East Asian crisis but find no significant return spillover from US to the Chinese A-share market.

Influential work by Bekaert and Harvey (1995) has attempted to date precisely the start of de facto financial integration of emerging stock markets. More recently, Ang and Bekaert (2002) have shown that Markov-switching models are well able to account for the asymmetric international correlation of equity returns. Characterising returns by two regimes: a normal one and a low-return-high-volatility market, they document that international correlations are larger in the latter regime than in the former. Regime-switching methods have also been used for the stock market by Huang (2000) for the beta in the CAPM, and Assoe (1998) for modelling the behaviour of returns in emerging markets.

The main focus of existing empirical work on China's financial integration is the A-share market. Indeed, while at first sight in a study of international integration it would be tempting to examine the market for B-shares, which is accessible for purchase to foreign investors, the thinness of trading on the B-share market has been shown to be responsible for spurious autocorrelation (Greonewold at al. 2001).

III. Methodology: Markov-switching vector error correction model

In a regime-switching model of returns some or all parameters depend on an underlying unobservable stochastic variable s_t, which aims at representing the phases of the returns

regimes (Krolzig, 2001). This approach enables us to assign probabilities to the occurrence of the different regimes. In its most popular version, which we will use here, such a model assumes that the process s_t is a first-order Markov process (Hamilton, 1994). The univariate Markov-switching procedure suggested by Hamilton was extended to non-stationary systems, in particular to vector error correction systems, by Krolzig (1997).

From Engle and Granger (1987) we know that with q_t a M-dimensional vector time series process, $q_t = [q_{1t}, q_{2t}, \ldots, q_{Mt}]'$, where q_t (denoting stock indices) is nonstationary but first-difference stationary, given $q_t \sim I(1)$, there could be up to M-1 cointegrating relationships, which represent the long run of the system. Krolzig (1997) showed that, if cointegration is present, it is possible to write the joint dynamics of the vector time series process as a Markov-switching vector error correction model (MS-VECM), as follows:

$$(1)\ \Delta p_t = \nu(s_t) + \Sigma_{d=1}^{n-1} \Gamma_d(s_t)\, \Delta q_{t-d} + \chi(s_t)\, \Pi\, q_{t-1} + e_t$$

with $\nu(s_t) = [\nu_1(s_t), \nu_2(s_t), \ldots, \nu_M(s_t)]'$ a M-dimensional column vector of regime-dependent intercept terms. The Γ_d's are M x M matrices of parameters, $\Pi = \alpha\beta'$ is the M x r cointegrating matrix with r the number of cointegrating relationships; ε_t a M-dimensional vector of error terms, $\varepsilon_t \sim NIID(0, \Sigma_\varepsilon(s_t))$. We assumes that the regime-generating process is an ergodic Markov chain with a finite number of states $s_t \in \{1,\ldots., K\}$ governed by the transition probabilities $p_{ij} = Pr(s_{t+1}) = j \mid (s_t) = i)$, and $\Sigma_{d=1}^{n-1} p_{ij} = 1$ for all i,j, with $j \in \{1,\ldots., K\}$.

We consider a Markov-switching Intercept-Autoregressive-Heteroskedastic VECM or MSIAH-VECM, in which speeds of adjustment to equilibrium can vary between regimes:

$$(2)\ \Delta p_t = \nu(s_t) + \Sigma\, \Gamma_d(s_t)\, \Delta p_{t-d} + D_t\, \Pi^1(s_t)\, p_{t-1} + (1-D_t)\, \Pi^2(s_t)\, p_{t-1} + e_t$$

with D_t a dummy variable taking a value of one over the period of validity of the first cointegrating relationship, and zero when the second cointegrating relationship is valid. We thus consider to what extent the long run relationship may be changing over time rather than being necessarily invariant over the full period of estimation. In other words, the β coefficients may differ over time. The intercept $\nu(s_t)$ also switches between states. With Markov-switching heteroscedasticity, the variance of errors can also differ between regimes. Finally, the autoregressive parameters (γ's), and the error correction coefficient (α's) are also

allowed to switch between states. We use likelihood ratio tests to check that such sources of switching are statistically acceptable[1].

The estimation of such a MS-VECM is legitimately implemented with a two-stage procedure (Saikonen and Luukonen, 1997). In a first step the Johansen (1998, 1991) maximum likelihood multiple cointegration method is used to determine the number of cointegrating relationships as well as to estimate the long run parameters.

(3) $\Delta p_t = \nu + \Sigma \Gamma_d \Delta p_{t-d} + \Pi p_{t-1} + e_t$

We pay particular attention to the instability of the long run relationships and in particular to the changes in long run attractors between periods.

In a second step, an expected maximization algorithm for maximum likelihood estimation is used to obtain estimates of the parameters in the Markov-switching model (Krolzig, 1997). For a given parametric specification of the model, (constant) probabilities are assigned to the unobserved return regimes conditional on the available information set which constitute an optimal inference on the latent state of the economy. We thus obtain the probability of staying in a given regime when starting from that regime, as well as the probability of shifting to another regime. The classification of regimes and the dating of returns periods imply that every observation in the sample is assigned to one of the regimes. The rule followed to assign an observation at time t to a specific regime depends on the highest smoothed probability. The smoothed probability of being in a given regime is computed by using all the observations in the sample. We assign an observation to a specific regime when the smoothed probability of being in that regime is higher than one half.

IV. Data and long run relationships between A-share market and global or regional markets.

Data and descriptive statistics

Given the time difference between East Asia and the United States, no contemporaneous quotes exist at the daily frequency. We thus use data at the weekly frequency (middle of the week), for the period from the third week of April 1993 through the

[1] When testing the Markov-switching model against the linear alternative or a *k* regime model against an (*k*-1) regime model, standard distribution theory does not apply (Davies, 1977) since a nuisance parameter (i.e. the transition probabilities) is not identified under the null hypothesis. The test proposed by Hansen (1992) and Garcia (1998) is conservative, tending to be under-sized and of low power. Ang and Bekaert (1998) conducted Monte Carlo experiments which imply that the true underlying distribution may be approximated by a $\chi^2(q)$ distribution, with q the sum of the linear restrictions and nuisance parameters.

third week of October 2004. We consider the logarithm of the Shanghai A-share index, of the Hang Seng Index, and of the Standard and Poor's 500. The source of the data is Datastream. We convert all indices into U.S. dollars.

We tested for the null of stationarity of the share price indices using two different unit root tests with non-stationarity as the null: Augmented Dickey Fuller (ADF), Philipps-Perron (PP) tests, as well as the Kiatkowski Phillips-Schmidt and Shin (KPSS) test which has the stationarity as null hypothesis. In all cases (table 1) we found that each of the three stock prices is integrated of order one. In the Shanghai case there is evidence supporting the presence of a deterministic trend.

Table 1: Unit root tests

Specification		With Intercept			With intercept	& trend
Type of test	ADF	PP	KPSS	ADF	PP	KPSS
Shanghai				-3.06	-2.94	0.29
[p-value] for signif trend				[0.01]*	[0.01]*	[0.00]**
Hang-Seng	-2.78	-2.92	0.77	-2.70	-2.89	0.28
[p-value] for signif trend				[0.70]	[0.70]	[0.00]**
S&P	-1.67	-1.69	2.28	-0.83	-0.83	0.72
[p-value] for signif trend				[0.76]	[0.98]	[0.00]**
(cv 1%)	(-3.43)	(-3.43)	(0.74)	(-3.97)	(-3.97)	(0.21)
(cv 5%)	(-2.86)	(-2.86)	(0.46)	(-3.41)	(-3.41)	(0.14)

Sample: Third week of April 1993 through the second week of October 2004

For the Shanghai A-share market index, the presence of outliers is very troublesome. Indeed, given the widespread interference by the authorities in the day to day behavior of the market, returns can exceed plus (or minus) 30% for one or two weeks at a time (see Girardin and Liu, 2005 on the distorting role of such outliers). Oral intervention by the stock market regulator is often responsible for such wide movements. Descriptive statistics on the raw data are presented in Table 2.

Table 2: Descriptive statistics for raw and adjusted returns

	Shanghai raw	adjusted	Hang Seng raw	adjusted	S&P raw	adjusted
Mean	-0.0197	-0.0197	0.123	0.123	0.15	0.15
Max	36.1	17.6	13.2	10.7	12.3	7.88
Min	-41.1	-14.3	-14.2	-14.2	-12.1	-6.61
σ	5.28	4.39	3.62	3.49	2.35	2.18
Skewness	0.14	0.375	-0.40	-0.28	-0.13	-0.19
Kurtosis	18.2	4.99	4.19	3.54	6.30	3.53

Visual inspection of figure 1a, confirms the size of outliers in returns. Using raw data for cointegration tests leads to sharp violations of normality assumptions. Even though the results may be qualitatively similar, they are not statistically robust. In his work on the way to conduct cointegration analysis in the presence of outliers, Nielsen (2004), on the basis of Monte Carlo simulations, shows that the traditional use of additive outliers can seriously distort inference on the cointegration rank and the long run parameters. Approximating additive outliers with combinations of unrestricted dummies is not a solution. The potential distortion from the use of an incorrect dummy specification is much larger than the distortion from ignoring outliers. However, using a linear interpolation, replacing an outlying observation by the simple average of the (level of) observations which precede and follow it, yields results which are very close to those obtained by more sophisticated procedures. We thus followed this procedure for each index, selecting outliers on the basis of large standardized residuals in the systems estimated. The result is called "adjusted" data in the following (see Table 2 for descriptive statistics on such data). We also ran all cointegrating and Markov-switching estimations with both additive outliers and combinations of unrestricted dummies and obtained qualitatively similar results, so we do not report the latter. However, we report in the Appendices the results with raw data.

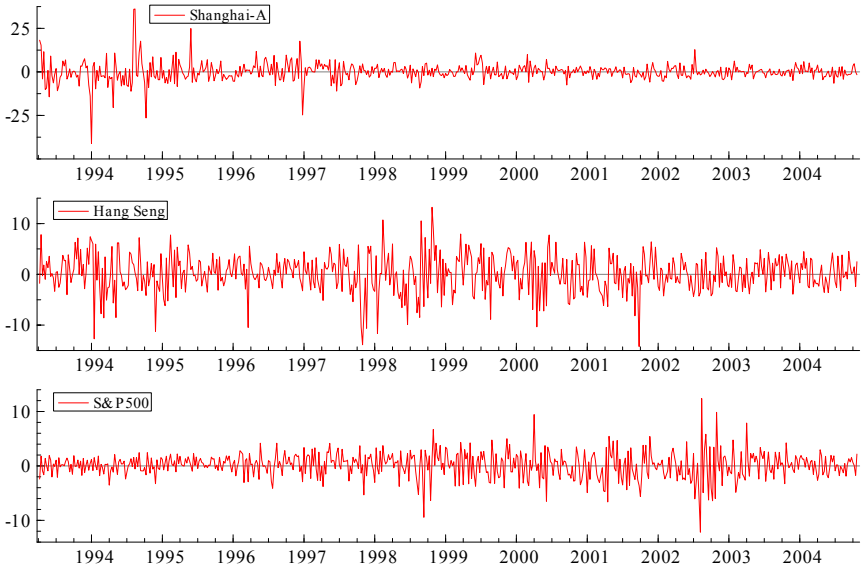


Figure 1a to 1c: Weekly returns in dollars for Shanghai-A, Hang Seng and S&P500: raw data

Cointegration tests

We first ran a finite order VAR model in level for the Shanghai A-share market and the Hang Seng, and S&P500 indices. All information criteria, (Akaike, Hannah-Quinn and Schwartz) conclude that the optimum lag is three. The same was true in bivariate systems. We then implemented the maximum likelihood procedure developed by Johansen (1988, 1991) to detect cointegrating relationships in a vector autoregressive system with an intercept and a trend. We rely on the Likelihood ratio trace test with the small sample critical values computed by Reimers (1992). The null hypothesis is that there are no more than r cointegrating vectors, with r between zero and p, for a vector q of variables: (q_i, i=1,.., p).

Table 3: Cointegration test (three lags): Weekly adjusted data in dollars.

	Trace [p-value]		
Rank r>=	0	1	2
(93-04)			
HK,S&P,T	33.0 [0.33]	14.4 [0.62]	4.90 [0.61]
S&P,T	21.7 [0.15]	6.23 [0.44]	
HK,T	18.1 [0.33]	5.53 [0.53]	
(93-97)			
HK,S&P,T	50.7 [0.00]**	24.5 [0.07]	11.0 [0.09]
S&P, T	32.1 [0.00]**	7.07 [0.35	
HK,T	18.5 [0.31	6.08 [0.46]	
(97-04)			
HK,S&P,T	46.5 [0.01]*	17.8 [0.35]	5.6 [0.52]
S&P,T	15.5 [0.53]	4.97 [0.61	
S&P	14.6 [0.25]	4.20 [0.39]	
HK,T	21.0 [0.18]	4.00 [0.74]	
HK	20.4 [0.04]*	3.35 [0.53]	

3 lags
** 1% level.; *5% level. [p-value]
SH= Shanghai-A, HK= Hang Seng, S&P= Standard and Poor's 500. T= Trend.

We considered the sample from the third week of April 1993 to the third week of October 2004. The results presented in Table 3, imply that , with a time trend, there is no evidence of cointegration between the Shanghai A-share market index, the Hang Seng and S&P indices. Neither is there evidence, in table 3, of bivariate cointegration with the Standard and Poor's or with the Hang Seng. However, when we spit the period around late September 1997 (on the basis of recursive cointegration tests), we find cointegration in the trivariate system over both sub-periods. Such relationships involve a time trend before but not after the East Asian crisis, as detected by exclusion tests (Table 4). Such tests also provide strong evidence that the Hang Seng is excluded from the long run relationship prior to the crisis, and (weaker) evidence that the S&P is excluded after that crisis. Bivariate cointegration tests fully confirm these results (Table 3). Shanghai is cointegrated with New York prior to the crisis, with a deterministic time trend, and with Hong Kong after the crisis, with no trend.

Overall, from the start of active trading in the latter in April 1993 through September 1997, the New York market represented the long run reference for the Shanghai market. However, since the East Asian crisis the A-share market has developed a substitute long run relationship with the Hong Kong market. We will consider in the next section to what extent such evidence in favour of an evolving long run anchor for the Chinese stock market can be supported by regime switching in the error correction process at the time of the East Asian crisis.

Table 4: Exclusion and tests: Weekly data in dollars

(β=0)	A-market	S&P 500	Hang Seng	Trend
93-97	2.83 [0.08]	4.57 [0.03]*	0.00 [0.93]	8.59 [0.00]**
97-04	8.89 [0.00]**	3.34 [0.07]	11.4 [0.00]**	

Likelihood ratio [p-value]; ** 1% level.; *5% level.
Exclusion of trend rejected for 93-97 with S&P, LR= 11.1[0.00]**

V. Markov-switching vector error correction.

In order to examine the evidence in favor of regime-switching error correction, we use a vector error-correction Markov-switching model. Our basic specification includes a trivariate system with returns in Shanghai-A, Hang Seng and the S&P. We implement specifications tests in three stages: linearity, nature of regime switching and the role of the error correction terms.

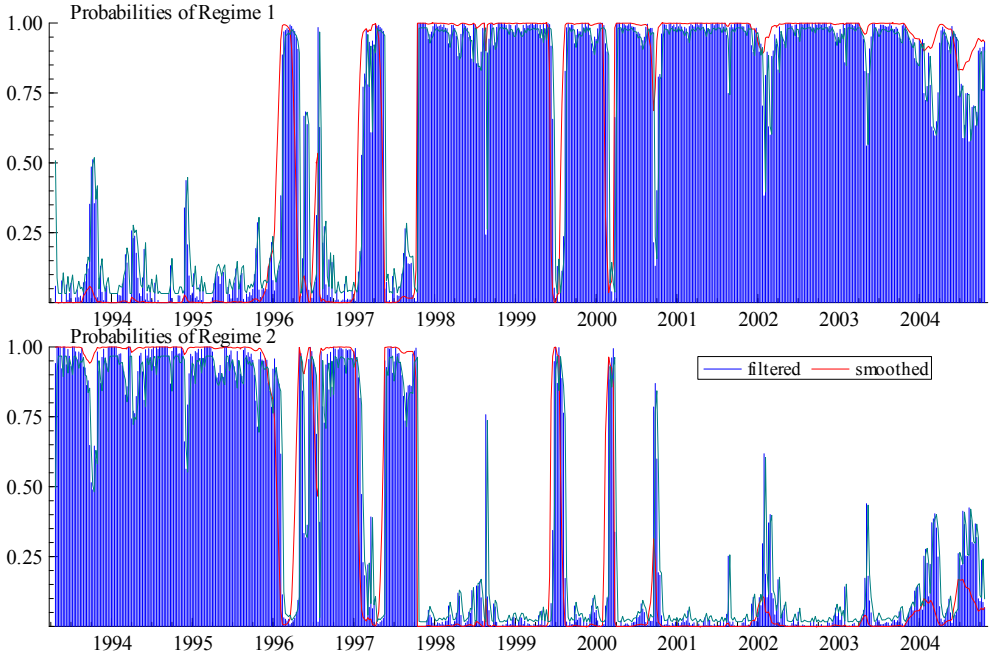
Table 5: Markov-switching error correction model: specification search.*

	Log Likelihood	2 lags	Linearity	MSIH	no ECM	no ECM1	no ECM2
MSIAH 2 regimes	-4391.6	3.0 [0.99]	261.8 [0.00]**	38.0 [0.00]**	80.0 [0.00]**	14.0 [0.03]*	20.4 [0.00]**

Markov Switching (MS) models have a switching intercept (I), switching variance (H) and switching autoregressive coefficients (A). LR=Likelihood Ratio; [p-value]. 1 lag.
Weekly adjusted data : 1993(15)-2004(43)

A likelihood ratio test concludes in favor of a two-regime model against the linear model (table 5, column 3). This shows that previous work using time-invariant linear models missed an important dimension. In the next step, it is important to pinpoint the source of regime switching within the two regime model. As reported in Table 5, column 4, the joint null hypothesis of no change in the autoregressive coefficients and of the error correction terms (MSIH) is rejected at the one per cent level in favor of the switching autoregressive model (MSIAH). Thus regime switching is not confined to the intercept and the variance. The hypothesis of insignificance of the error correction terms is rejected at the 1% level (Table 5, col. 5). Individually, each error correction term also appears significant (Col. 6 and 7, Table 5).

Figure 2: Probabilities of regimes in the Markov-switching vector error correction model.



Figures 3 plots the estimated filtered (using information up to the previous period) and smoothed (using the whole sample) probabilities for the two regimes for the Markov-switching asymmetric error correction models. The timing of the regimes implies a permanent change in the early fall 1997, i.e. at the time of the East Asian crisis. However regime One arose already briefly before the crisis, in 1996 and again in early 1997.

Table 6: Estimated coefficients of the Markov-switching vector error correction model: Adjusted data, Shanghai equation

Regime	1	2
Intercept	0.105	-0.56
	(0.72)	(1.32)
ΔSH_{-1}	0.039	0.19
	(0.69)	(2.92)
ΔHS_{-1}	-0.061	0.058
	(1.37)	(0.41)
$\Delta S\&P_{-1}$	0.12	0.672
	(1.83)	(2.15)
$[p-a_1-\beta_1(S\&P1)]_{-1}$	0.009	-0.057
	(0.41)	(2.62)
$[p-a_2-\beta_2(HS2)]_{-1}$	-0.022	-0.09
	(2.96)	(1.72)
σ	2.63	5.84

(t statistic); sample: 1993(15)- 2004(43)
[p-a- β(S&P1)] = error correction term with S&P over 1993(15)-1997(39)
[p-a- β(HS2)] = error correction term with Hang Seng over 1997(40)-2004(43)

As shown in Table 6, the post-crisis regime (regime One) corresponds to zero average returns and low volatility, while the pre-crisis regime (regime Two) has negative returns with high volatility. All regimes are very long-lasting and stable, since the probability of moving out of a given regime is never larger than 0.03. We checked that the regime probabilities are not contingent on the use of our split error correction terms. Indeed when excluding the latter, in a VAR in first difference, the probabilities of regimes are identical to the case when they are present. We thus do not report them here.

In the pre-crisis regime, with the error correction term corresponding to the long run relationship with the S&P the speed of adjustment was larger than that corresponding to the error correction with respect to the Hang Seng index in the post-crisis regime. Besides, prior to the crisis, Shanghai returns had a significant autoregressive component of order one (a property which can be induced by the use of adjusted data), and were influenced by lagged New York returns. These two effects all but vanished thereafter.

In order to check that the use of adjusted data does not distort the results through data mining, we rerun an identical specification with raw data. The underlying cointegrating

relationships have the same specification as above, but no treatment is implemented for outliers. The results, given in Appendix 1, are very close not only qualitatively but also quantitatively to the results of the estimation with adjusted data. The price to pay is apparent when comparing the behavior of residuals in the models with the two data sets, as shown in Appendix 2.

VI. Explaining the evolving international integration of Chinese equity markets.

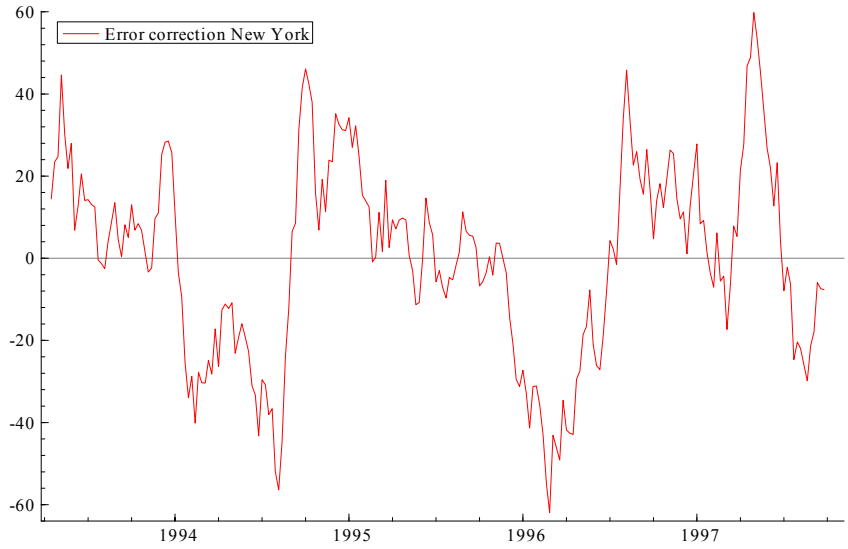
Three results of our MS-VECM analysis deserve explanations: i) the anchor represented by the New York market over the pre-crisis period; ii) the full disappearance of such an attractor over the subsequent period; and iii) the emergence of the Hang Seng index as a new anchor after the East Asian crisis.

During the pre-crisis regime only error correction with New York was at work (see figure 3a). Such significant error correction must have been due to information flows. This regime was characterised by high volatility in returns on China's stock market and any disturbance could easily be amplified, especially since, during this period, investors were mostly naïve individuals. Institutional investors indeed played a very limited role. These individual investors took the US stock market as a model and took into account what happened in that market. They must particularly have noticed that the index was at a relatively too low level (for example in early 1996).

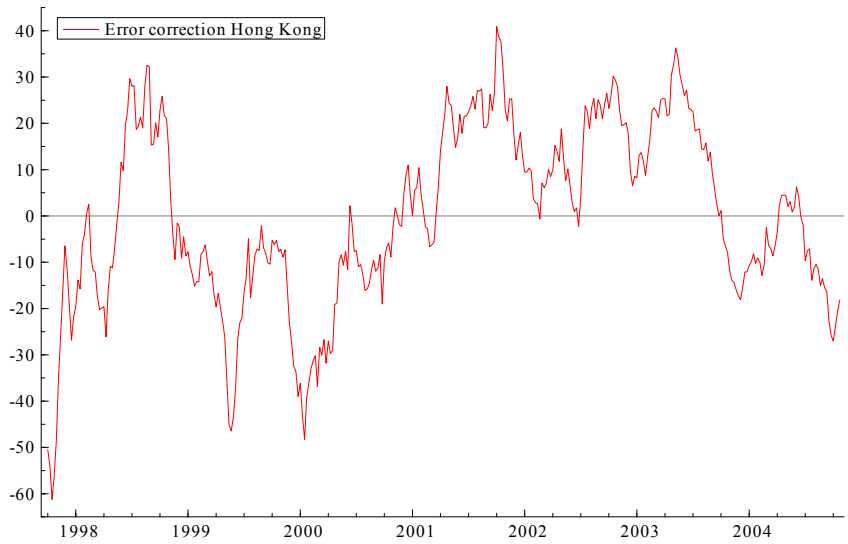
The vanishing influence of the New York anchor after the crisis can be explained on different grounds. Indeed, after 1997 when the US internet fever started driving US stock prices to ever higher levels, the Chinese market was depressed and remained so until early 1999. One of the reasons is that the internet fever spilled over to China only from the middle of 1998 onwards. The penetration of ICTs in China was a little slow initially. The (late) catching up the bubble phase was extremely rapid and concentrated over a short period. Trading volume doubled in 2000 compared to 1999. When the US bubble burst in March 2000, the bubble in China was still developing. By the end of 2000, Chinese regulators started being concerned with this overvaluation of the market and, in 2001, attempted to cool down the fever. The authorities started controlling illegal trading, sped up IPOs, and discussed making more SOE shares tradables.

Figure 3: Eror correction term:

a. With respect to S&P: April 1993-September 1997



b. With respect to Hang Seng: October 1997-October 2004



The new role played by the Hang Seng index (see figure 3b for the error correction term) can be understood to the extent that since the middle of 2001 with the sharp fall in the market, regulators and investors tried to find a new anchor for the market. They both compared China's P/E ratios with Hong Kong ones. This built up the convergence process between Chinese stock prices and regional prices. At the beginning of this period, in early 2001, the Chinese regulators had introduced price-depressing measures deliberately in order to reduce P/E ratios towards international levels, on top of making the bubble burst. The

convergence process then initiated and subsequently deepened was thus partly the outcome of a policy decision. Since mid-2001 the A-share market started to adjust downwards and the P/E ratio has been reduced from 50 or 60 to 20-30.

Since middle 2001 the A-share market in China kept being a bear market. More and more market participants tried to connect this to Chinese stock market converging to international stock markets, leading to a continuous fall in the A-share price. People paid more and more attention on comparing prices of mainland companies listed in Shanghai (A-shares) and in Hong Kong (H- shares). For example in 2004 A-share prices were often 40% higher than H-Share, some even were in excess of 100%. Since more and more large Chinese companies have multiple listing in HK and domestic markets, investors start to worry that A-share prices and international prices will have to soon.

Chinese observers of the international integration of the A-share markets are mostly focused on low frequency comparisons of Price-Earnings ratios. In the Summer 2004 one of the most influential Chinese financial journals (Security Market Weekly) had a special report on this topic (Chen and Kai, 2004), which also only looked at P/E ratios. Most of their conclusions were that the P/E ratio is currently too high for A-share markets, which implies that share prices in China are comparatively higher than the international standard. Similarly, when Chinese firms are listed on foreign markets on top of Chinese ones, the price of their shares is always much higher inside than outside China. For such observers, what matters for the international convergence of stock markets is not always capital flows. The thinking of investors is more important than capital flows. Chinese investors have reduced their thought difference with international investors dramatically in the past five years. Even with strong capital controls the internationalisation of the investment strategy of Chinese investors will strongly support the downward convergence of A-share prices to international levels.

The convergence of share prices between Shanghai and Hong Kong still has a long way to go. Indeed, at the end of our sample P/E ratios were still close to 30 in Shanghai as opposed to 15 for the Hang Seng. The convergence of A-shares H-shares will not be in one direction (a fall in A-share prices), but should be two-directional because H-shares are undervalued and most Chinese investors believe this is the case. Since 2003 more and more Chinese mainland companies have been listed in HK stock market, but due lack of transparency and credibility of such companies, international investors cannot be confident enough to buy their shares, leading to a low P/E ratio on the H-share market. But Chinese

mainland investors know Chinese listed companies better than international investors. When Chinese investors will be allowed to buy H-shares, their price will rise.

VII. Conclusion.

Existing empirical work (using data up to 2001) has concluded that the Chinese A-share market is not integrated, either in the short or in the long run, with regional or international stock markets. Previous work also found no significant evidence of a break in such relationships at the time of the East Asian crisis. Using the longest available sample of active trading in the Shanghai A-share market (April 1993 through October 2004), with weekly data in dollar terms, the present paper provides contrary evidence on both fronts.

Indeed, our cointegration tests document the existence of a long run relationship between the Shanghai A-share market index and either the Standard and Poor's 500 or the Hang Seng index. Neither relationship is valid for the overall period. While the former prevailed prior to the East Asian crisis, after that crisis a new long run relationship arose for the Shanghai market with the Hang Seng index.

We used a Markov-switching vector error correction model for A-share returns to study how dynamic links have been regime dependent. Prior to the East Asian crisis negative deviations form the long run relationship with New York led to substantial and significant error correction, while lagged US returns had a substantial impact on Shanghai returns, while subsequently the only dynamic effect involved correction of deviations from the long run link with Hong Kong, with a lower speed of adjustment.

Several reasons can explain the differences between our results, supporting in turn international and regional integration of the Chinese stock market and previous results against it. Our use of weekly data as opposed to daily data often favoured by previous work is not responsible for the difference. Daily data should not be used for cointegration tests between Chinese and U.S. stock returns because of the absence of synchronous trading. The lack of finding of cointegration in previous work may be due to the neglect of deterministic trends and the lack of proper treatment of outliers. Such outliers are a major source of bias since they are extremely large in an emerging stock market as peculiar as China's. Another reason for the difference of finding is linked to the necessity of allowing for the possibility of regime switching in the dynamic links.

There are favourable prospects for the opening of China's financial market. Currently a Qualified Domestic Institutional Investors' (QDII) scheme, which would allow large

domestic institutional investors to invest abroad (mainly in the Hong Kong Market), is actively discussed. Similarly Chinese Depository Receipts, allowing H-shares to be traded in domestic markets, are also likely to be introduced soon. All these would represent potential further channels for connexions with international markets. The relationships with the Hong Kong market should be strengthened in as much as there is a potential for the rise of share prices on that market while prices in Shanghai continue their decline.

Appendix I

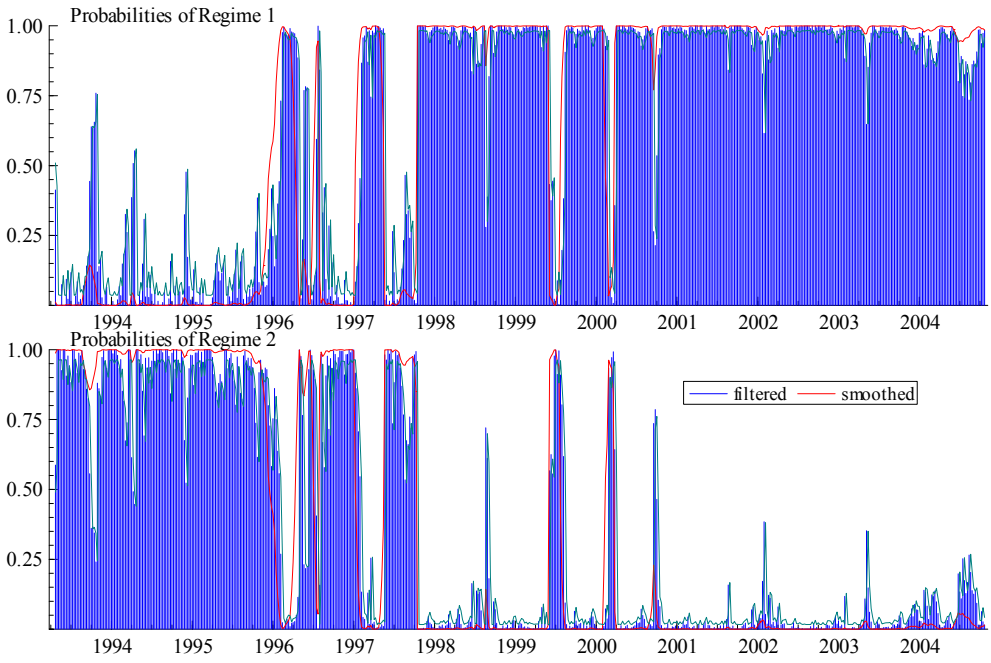
Markov-switching vector error-correction model with raw data

Table A1: Estimated coefficients of the Markov-switching vector error correction model: raw data, Shanghai equation

Regime	1	2
Intercept	0.07	-0.72
	(0.50)	(1.24)
ΔSH_{-1}	0.006	0.06
	(0.11)	(0.90)
ΔHS_{-1}	-0.074	0.062
	(1.69)	(0.35)
$\Delta S\&P_{-1}$	0.153	0.787
	(2.56)	(1.83)
$[p\text{-}a_1\text{-}\beta_1(S\&P1)]_{-1}$	0.03	-0.135
	(1.93)	(1.86)
$[p\text{-}a_2\text{-}\beta_2(HS2)]_{-1}$	-0.021	-0.054
	(2.69)	(2.30)
σ	2.67	5.84

(t statistic); sample: 1993(15)- 2004(43)
[p-a- β(S&P)] = error correction term with S&P (with deterministic trend) over 1993(15)-1997(39)
[p-a- β(HS2)] = error correction term with Hang Seng over 1997(40)-2004(43)

Figure A1: Probabilities of regime, MS-VECM with raw data



Appendix 2:
Standardised residuals with raw and adjusted data
Shanghai equation

Figure A.2.1.Raw data

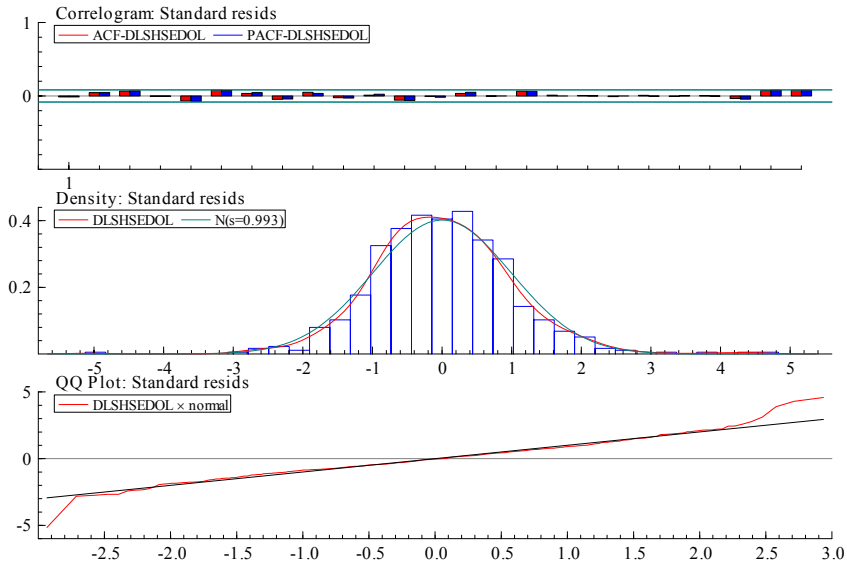
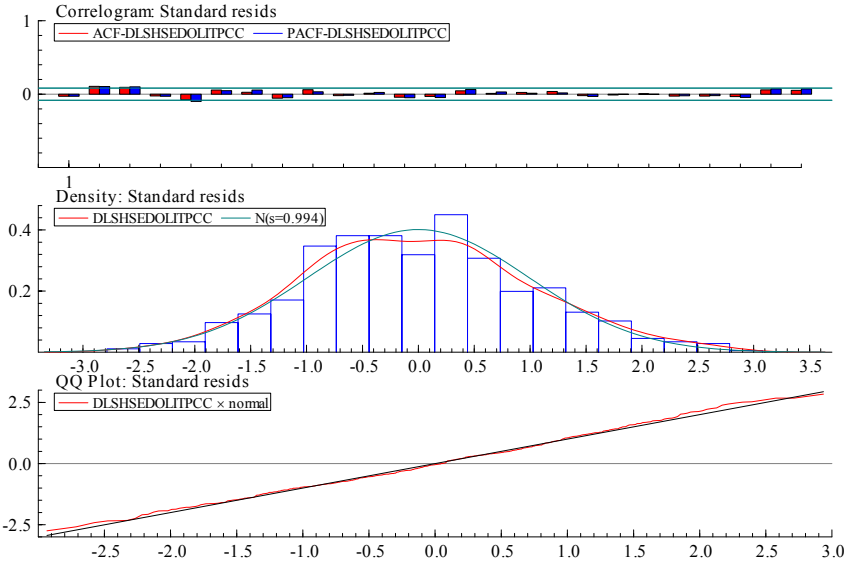


Figure A.2.2. Adjusted data



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