

The financial integration of China: New evidence on temporally aggregated data for the A-share market

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Eric GIRARDIN & Zhenya LIU

Abstract

In spite of high trade openness, existing empirical work, using daily data, has not found any evidence of international financial integration of China. In this paper we examine to what extent the Chinese A-share market, *de jure* protected from foreign influences by capital controls, is actually integrated with global or regional markets. We study a long sample (October 1992 through March 2005) of active trading, within the framework of a regime-switching error correction model. We confirm the role of temporal aggregation in cointegration tests. With daily or mid-week closing prices, we do not find any long run relationship with either the New York or the Hong Kong market, thus replicating previous findings. However, the use of weekly averaged prices implies that, up to late 1996, the Shanghai A-share market index was cointegrated with the S&P500. Subsequently, this relationship broke down and a long run relationship with the Hang Seng index gradually arose. Information flows, as well as the prospects of *de jure* financial opening, and the growing awareness of valuation concepts among Chinese domestic investors, in the presence of multiple listing of Mainland firms, help explain the evidence of financial integration in spite of capital controls.

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Keywords: China's A-share market, Markov-switching ECM, temporal aggregation, international financial integration

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I. Introduction

Resisting the East Asian wave of external financial liberalisation in the 1990s, China has retained capital controls, preventing foreigners to access A-share markets and residents to access foreign stock markets (Lardy, 1998). However, since the end of 1999, internationalisation has become one of the big themes of Chinese stock market reform. Going beyond WTO commitments, which do not concern liberalisation of portfolio flows, China allowed foreign investors 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' scheme. *De jure* financial openness has thus made some progress, leading to a presumption that the integration of China's capital markets would have increased in the new millennium.

However, existing work on emerging markets show that, in the presence of high trade openness, external *de jure* financial openness is neither sufficient nor necessary for the *de facto* openness of domestic capital markets (Aizenman, 2003). The degree of trade openness in China has certainly reached high levels and hidden capital (out)flows have been very large, especially in the second half of the nineties (Gunther, 2004), accumulating to more than the stock of inward FDI.

Still a consensus view among empirical researchers is that the Chinese stock market has been mostly protected from the influence of foreign stock markets. Existing work, using widely different samples (always ending in 2002 at the latest) with high frequency (daily) data, does not find any evidence of long run relationships between A-share indices and foreign stock market indices. Neither does this literature document the emergence of such relationships at the time of the East Asian crisis.

In this paper we re-examine the issue of the existence of long run relationships between Shanghai's A-share market and international stock markets, over a long sample of active trading (October 1992-March 2005). 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 one fourth of the Hang Seng index lists Mainland Chinese firms. We use a regime switching error correction model which is more flexible than other techniques in detecting gradual changes in long run anchors. This enables us to determine whether there was a change in regime in

international financial integration of China at the time of the East Asian crisis, or whether seeds of change were already present before the latter.

We proceed in two steps. Firstly, using daily or mid-week closing prices, we replicate, on our longer sample and with this different methodology, the results of previous work on the absence of cointegration. Secondly, in line with existing work supporting the use of temporally aggregated data when testing for the presence of long run relationships, the study of weekly averaged data leads us to provide fresh evidence of evolving cointegration. There was a change in long run anchor for the Shanghai A-share market from a global one, the Standard and Poor's 500, up to 1996, to a regional one, the Hang Seng index, subsequently. This enables us, in particular, to explain what is a puzzle to millions of individual Chinese investors, and many observers, i.e. the everlasting bear phase of the secondary market since 2001.

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 arguments supporting temporal aggregation when testing for cointegration will be presented and the methodology for Markov-switching error correction models will be introduced. Section four will present the data, and provide the results of the estimation of the error correction model both replicating the results of previous work and providing new evidence. In section five we explain the background reasons for the evolving international integration of China's stock market. Section six concludes.

II. A consensus in existing literature: The A-market is not internationally integrated

Existing theoretical work has provided rationales and empirical work has offered supporting 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. On the theoretical side, Aizenman (2003) showed that, with high trade openness, *de jure* protection of a domestic financial market through capital controls does not prevent *de facto* financial openness. On a descriptive side, using measures of *de facto* financial openness based on gross capital flows as a share of GDP, Lane and Milesi-Feretti (2005, p.9) find that China is an example of a "group of countries [for which] gross inflows and gross outflows have both been large and roughly similar in magnitude, reflecting increased financial integration with the world economy". China has experienced large capital inflows, mainly through FDI, and

outflows, but also significant net foreign asset accumulation, mostly as foreign exchange reserves. Potential channels for hidden capital flows to and from China are important. By itself, the evidence on the huge scale of round-tripping behind FDI, which according to Xsiao (2005) has a mean value of 40% of recorded FDI, shows that such hidden flows do take place. Indeed, round tripping FDI assumes that capital flight has already taken place. Estimates by Gunther (2004) of the extent of capital flight out of China, using all available methods and making the necessary adjustments, recently updated by Ljungwall and Wang (2004), imply a sharp rise in capital flight in the second half of the nineties, to reach a peak of 80 billion dollars in 1998, falling to 64 billion in 2003, accumulating since the mid-1980s to more than the net FDI stock. Round tripping in the form of portfolio investment also occurred on the inflow side, even before the opening of the Chinese B-share market to domestic residents in 2001, when a large number of the Chinese residents investing in B shares were using foreign borrowed passports and foreign bank accounts.

In their study of financial globalization, Prasad et al. (2003) have uncovered, for emerging market countries, both cases of 'financial integration without capital account liberalisation' and 'liberalisation without integration'. On the econometric side, for East Asian emerging market countries, excluding China, Phylaktis and Ravazzolo (2002) have vindicated this distinction on monthly data. Using Johansen's (1991) multiple cointegrating techniques, they find that the relaxation of ownership restrictions in the 1980s was 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 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 (Goenewold et al. 2001).

Many empirical studies of financial integration have examined the case of China, always with daily data. Huang et al (2000) consider the causal and long run relationships between China's market and the markets of Hong Kong, Japan, Taiwan or the US. Over the period 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. Using a vector autoregressive (VAR) model in first differences, over a recent, but short, period (September 2001 to December 2002), Hsiao et al. (2003) do not find either any causal relationship from the US market to the Chinese market, and with impulse response functions, show that a shock in US stock prices has no impact on Chinese stock prices.

Hatemi and Roca (2004) study 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 that the US stock market started influencing China after the East Asian crisis. Goenewold et al. (2004) examine the relationships between Mainland Chinese markets and the Hong Kong and Taiwanese markets over the October 1992 through November 2001 period. They confirm that Hong Kong is not cointegrated with China's A share markets, and, using a VAR in first differences, conclude that 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. Examining 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. Finally, Wang and Firth (2004), over the period November 1994 through September 2001, document lagged return spillovers from Hong Kong (but not New York) to Shanghai. However, they abstract from the issue of long run relationships to focus on volatility spillovers.

Markov-switching techniques were used by a number of studies for emerging markets, excluding China. Influential work by Bekaert and Harvey (1995) on time-varying financial integration used such techniques to date precisely the start of *de facto* financial integration. More recently, Ang and Bekaert (2002) have shown that such models are well able to account for the asymmetric international correlation of equity returns. Similar methods have also been used in a closed economy context. Huang (2000) employed it for the beta in the CAPM, and Assoe (1998) for modelling the behaviour of returns in emerging markets. In the case of China, these techniques were used by Girardin and Liu (2003), on weekly data, to detect speculative periods in the Shanghai A-market.

III. Methodology: Temporal aggregation and error correction models

All too often, temporal aggregation is not taken into account for the estimation of cointegrating relationships. It is even usual to assume that the use of high frequency data provides better results than low frequency data because of the bigger sample size allowed by the former. As seen in the previous section, all existing work on the international integration of the Chinese stock market has relied on such a premise to justify the use of daily data. However, recent work on the effects of temporal aggregation has shown that, in order to detect error correction and cointegration, we may be better off with temporally aggregated data. As shown by Granger (1990) and confirmed by Marcellino (1999), cointegration relations remain invariant to temporal aggregation. With data temporally aggregated by non-overlapping averaging, Rotger, Rossello and Caralt (2000) showed that the (OLS) estimator of a cointegrating vector is better, in terms of variance and bias, than the estimator using disaggregated data. Such a transformation of the data, through averaging, can be applied to flow variables as well as to stock variables since the cointegration space is not affected. Overall such temporal aggregation may improve the estimation of the long run coefficient. If there is cointegration, some adjustment (error correction) coefficient will remain non-zero regardless of the level of temporal aggregation. An illustration is provided for asset prices by Rajaguru and Abeysinghe (2003) in the case of the yen-deutsche mark exchange rate for daily, end-of-week and weekly averaged data. However, the magnitude of the error correction coefficient may be aggregation dependent (Marcellino, 1996, who provides an illustration using averaged data for the short term Canadian interest rate). Finally the reduction in sample size is not too worrying when working with asset prices and aggregating, for example, from daily to weekly data.

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 (Hamilton, 1994). 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, 1989). The univariate Markov-switching procedure suggested by Hamilton was extended to multivariate systems by Krolzig (1997). We examine here a particular case of that more general model with a one-step Markov-switching error correction model (MS-ECM), such as:

$$(1) \Delta p_t = v(s_t) + \sum_{d=1}^n \chi_d(s_t) \Delta p_{t-d} + \alpha(s_t) p_{t-1} + \xi(s_t) q_{t-1} + (\sigma)^{1/2}(s_t) \varepsilon_t$$

The χ_d 's are coefficients autoregressive distributed lags on stock returns, α is the error correction coefficient, p and q the logarithm of the domestic and foreign stock market price, and ($\beta = [-\xi/\alpha]$) the coefficient of the latter in the cointegrating relationship. We assume that the regime-generating process is an ergodic Markov chain with a finite number of states $s_t \in \{1, \dots, K\}$ governed by the transition probabilities $p_{ij} = \Pr(s_{t+1} = j | (s_t) = i)$, and $\sum_{d=1}^{n-1} p_{ij} = 1$ for all i, j , with $j \in \{1, \dots, K\}$.

All coefficients, plus the variance, are assumed to be regime-dependent. We thus examine a Markov-switching Intercept-Autoregressive-Heteroskedastic ECM or MSIAH-ECM. 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 between regimes. The intercept $v(s_t)$ also switches between states. With Markov-switching heteroskedasticity, the variance of errors can also differ between regimes ($[\sigma]^{1/2} [s_t]$). 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 acceptable¹.

An expected maximization algorithm for maximum likelihood estimation is used to obtain estimates of the parameters in the Markov-switching model (Hamilton, 1994). For a given parametric specification of the model, 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 (constant) 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. We assign an observation to a specific regime when the smoothed probability of being in that regime is higher than one half. The smoothed probability is computed by using all the observations in the sample.

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.

IV. Evidence of long run integration.

For the Shanghai A-share market index, the presence of outliers is troublesome. Indeed, given the widespread interference by the authorities in the day to day behaviour of the market, returns can exceed plus (or minus) 30% for one or two days 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. Using such data for cointegration tests leads to sharp violations of normality assumptions. Outliers can seriously distort inference on the cointegration rank and the long run parameters. Even though the results may be qualitatively similar, they are not statistically robust (Nielsen, 2004).

In order to try and replicate the results of previous work, we will first ignore outliers and check whether the absence of cointegration is robust to the use of a longer sample. We will thus use either daily or weekly (every Wednesday and every Friday) closing price data over the period from the first week of October 1992 through the third week of March 2005. We will then, to some extent, sidestep the problems linked to outliers by using, in a second stage, the weekly average of each (closing) stock market index, over the same sample (data plotted in Appendix I). We use 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 consider all indices in domestic currency. We checked that the results are qualitatively the same when indices are expressed in the same currency, which is not surprising given the stability of the RMB-dollar rate over most of the sample, so we do not report them.

We tested for the null of stationarity of the share price indices using a unit root test with non-stationarity as the null (Phillips-Perron (1988) test), as well as the Kiatkowski, Phillips, Schmidt and Shin (1992) test which has the stationarity as null hypothesis. In all cases (results not reported to save space) we found unsurprisingly, with and without a deterministic time trend, that each of the three stock prices is integrated of order one.

We use an error-correction Markov-switching model (equation 1), in which our basic specification is an equation for Shanghai-A returns with lags of such a variable, plus the (one period) lagged level of the Shanghai-A index, and either the Hang Seng or the S&P500 index. We proceed in two steps, first using daily, mid-week or end-of-week, closing prices, in order to try and replicate existing findings, and then employing weekly averaged indices. In both cases, we implement specifications tests in three stages: linearity, nature of regime switching and the role of the error correction terms.

Little or no evidence of integration with daily, mid-week or end-of- week data.

For both the model with mid-week and end-of week data, a likelihood ratio test concludes in favour of a two-regime model against the linear model (table 1, column 2). This shows that previous work using time-invariant linear models missed an important dimension. Columns 3 and 7 of Table 1 show that the two-regime model is rejected against the three-regime one. As reported in Table 1, columns 4 and 8, the hypothesis of insignificance of the error correction terms is rejected at the 1% level.

Table 1: Markov-switching error correction models, mid-week and end-of-week data: Specification search.

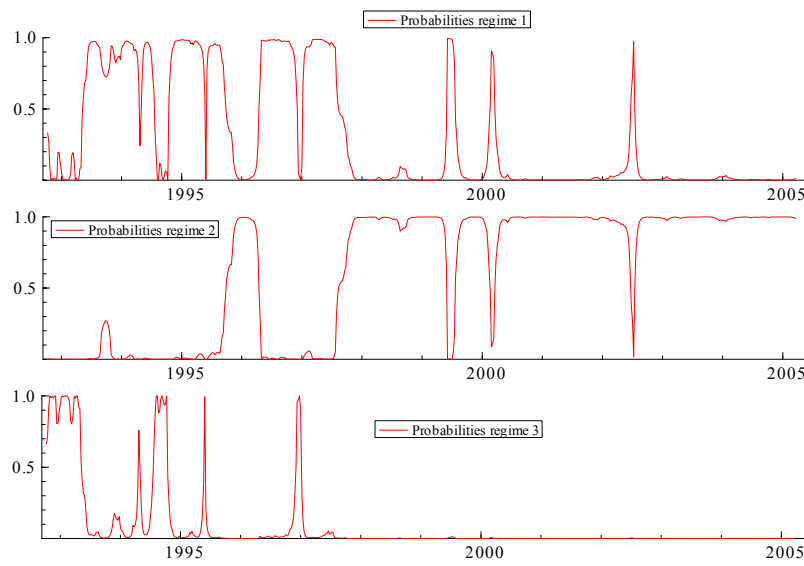
	Mid week				End of week			
	Log Likelihood	Linearity vs. 2 regimes	2 regimes	no ECM	Log Likelihood	Linearity vs. 2 regimes	2 regimes	no ECM
<i>With S&P500</i>	1173.5	343.8 [0.00]	144.2 [0.00]	27.2 [0.00]	1190.4	516.7 [0.00]	76.0 [0.00]	26.0 [0.00]
<i>With Hang Seng</i>	1169.60	352.0 [0.00]	63.6 [0.00]	24.8 [0.00]	1190.7	519.0 [0.00]	90.0 [0.00]	32.0 [0.00]

LR=Likelihood Ratio in all columns except col.1; [p-value]. Sample: 1992(47)-2005(12)

For mid-week data, Figure 1 plots the estimated smoothed probabilities for the three regimes for the Markov-switching error correction model involving the S&P index (the same plot is valid with the Hang Seng). Given that the probabilities for the models using end-of-week data are similar to figure 1, we do not report them. The timing of the regimes implies a permanent change after the mid nineties. Indeed, regime Two arose already in the last quarter of 1995 and the first four months of 1996, and then disappeared, until late September 1997. A third regime existed in late 1992 and the first half of 1993 but almost vanished thereafter.

As shown in Table 2a, for mid-week data, both with the Hang Seng and the S&P500, the third regime represented a period of high volatility in an immature market with very high average returns. The post-1996 regime (regime One) corresponds to zero average returns and low volatility, while the pre-1997 regime (regime Two) has negative returns with moderate volatility. All regimes are very long-lasting and stable, since the probability of moving out of a given regime is never larger than 0.16.

Figure 1: Probabilities of regimes in the MS-ECM: with S&P500 (Mid-week data)



Lagged S&P, or Hang Seng, returns were never found significant. With mid-week data, in both regime One and regime Two, error correction is never significant², either in the estimation with the S&P or the Hang Seng series. On the basis of these results, using closing mid-week prices, there is thus no evidence of cointegration between the Shanghai market and either the US or the Hong Kong market.

Table 2a: Estimated coefficients of the Markov-switching error correction models:
Mid-week data

Regime	with S&P			with Hang Seng		
	1	2	3	1	2	3
Intercept	-0.191 (1.50)	-0.066 (1.45)	0.526 (0.77)	-0.427 (2.39)	-0.089 (1.36)	2.32 (2.57)
ΔSH_{-1}	-0.103 (1.39)	0.016 (0.30)	0.378 (2.69)	-0.07 (0.98)	0.009 (0.16)	0.259 (1.84)
SH_{-1}	-0.047 (1.98)	-0.009 (1.36)	-0.126 (2.17)	0.022 (0.14)	-0.004 (0.62)	-0.232 (3.41)
X_{-1}	0.079 (3.69)	0.018 (2.03)	0.05 (0.50)	0.044 (1.88)	0.012 (1.56)	-0.078 (1.07)
σ	0.051	0.026	0.132	0.052	0.025	0.114
Probability of persistence	0.934	0.987	0.875	0.927	0.984	0.838

(t statistic); sample: 1992(47)- 2005(12); SH= Shanghai-A index, X= Standard and Poor's 500 (columns 1 to 3) or Hang Seng index (col. 4 to 6).

² In assessing the presence of a cointegrating relation, we follow Banerjee et al. (1993) who suggest that a bivariate error correction model with a t statistic greater than 3 for large samples provides strong evidence of a cointegrating relation (at the 5% level).

As reported in Table A.2 (Appendix II), similar estimates on daily data, over the same sample, provide no evidence of cointegration of the Shanghai market with a foreign market, while the three-regime decomposition is analogous to the one reported in figure 1. The use of end-of-week data (Table 2b) similarly does not provide any support for the presence of a long run relationship with New York, but offers some weak evidence of such a relationship with Hong Kong, in regime Two, with a coefficient close to one.

Table 2b: Estimated coefficients of the Markov-switching error correction models:
End-of-week data

Regime	with S&P			with Hang Seng		
	1	2	3	1	2	3
Intercept	-0.265 (2.73)	0.079 (1.69)	3.622 (3.37)	-0.382 (2.46)	-0.034 (0.51)	4.05 (2.98)
ΔSH_{-1}	0.046 (0.54)	0.078 (1.22)	-0.136 (1.07)	0.044 (0.05)	0.057 (0.84)	-0.117 (0.93)
SH_{-1}	-0.021 (1.03)	-0.013 (1.82)	-0.416 (5.38)	0.010 (0.89)	-0.018 (2.81)	-0.442 (4.57)
X_{-1}	0.061 (2.83)	0.002 (0.22)	-0.111 (0.81)	0.032 (1.66)	0.017 (2.16)	-0.105 (1.07)
σ	0.051	0.024	0.124	0.05	0.022	0.12
Probability of persistence	0.88	0.95	0.84	0.89	0.95	0.82

See note top Table 2a.

Integration with weekly averaged data.

As with the mid-week and end-of-week data, the use of weekly averaged data leads to the rejection of the two-regime model against the linear model (table 3, column 2), as well as against the three-regime model (col. 3). The hypothesis of insignificance of the error correction term is not accepted either.

Table 3: MS-ECM, weekly averaged data: specification search.

	Log Likelihood	Linearity vs. 2 regimes	2 regimes	no ECM
With S&P500	1173.5	401.2 [0.00]	62.2 [0.00]	16.2 [0.02]
With Hang Seng	1273.4	395.7 [0.00]	76.6 [0.00]	31.4 [0.00]

LR=Likelihood Ratio in all columns except col.1; [p-value]. 1992(47)-2005(12)

As shown in Figure 2, for the relationship with the Hang Seng, with the weekly averaged data, the start of regime Two occurs more than a year later than with the mid-week data, i.e. in January 1997. It then gives room again temporarily to regime One from late April to early October 1997. Subsequently regime One creeps back on four occasions: for one

month in the late Summer 1998, two and a half months in the Summer of the following year, from mid February to late March 2000 (around the New York crash), and for three weeks in late June-early July 2002. The probabilities with the S&P500 are almost identical, and not reported.

Figure 2: Probabilities of regimes in the MS-ECM: with Hang Seng (Weekly averaged data)

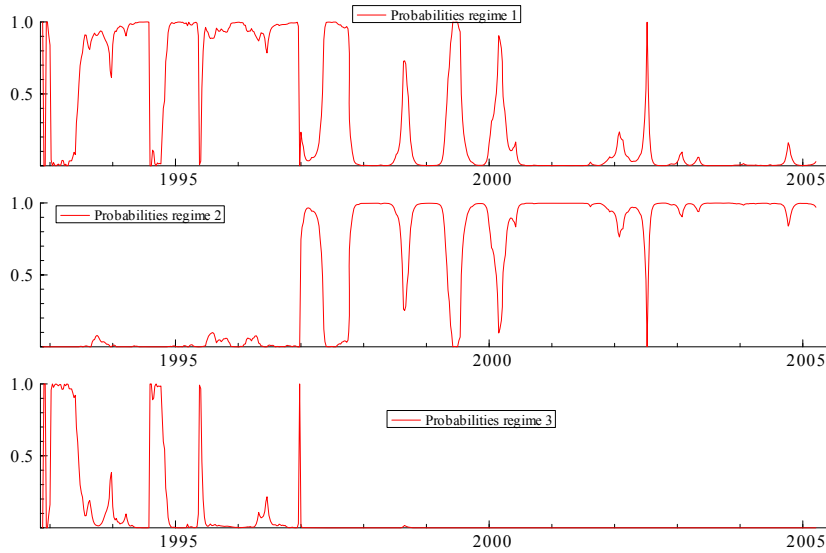


Table 4 provides very different results from those contained in Table 2. Indeed, in regime One, error correction with the S&p500 is now significant, with an adjustment coefficient close to $[-0.05]$, and a long run coefficient equal to $[1.39]$. By contrast, in regime Two, the error correction with the Hang Seng is activated significantly. However, both the error correction coefficient (in absolute value) and the long run coefficient are twice smaller than was the case with the S&P in regime One. Lagged foreign returns never proved significant and were accordingly dropped.

Table 4: Estimated coefficients of the Markov-switching error correction models:
Weekly averaged data

Regime	with S&P			with Hang Seng		
	1	2	3	1	2	3
Intercept	-0.111 (2.14)	0.217 (4.39)	0.606 (0.89)	-0.288 (2.27)	0.001 (0.01)	4.73 (6.29)
ΔSH_{-1}	0.106 (1.65)	0.259 (4.16)	0.411 (3.35)	0.09 (1.51)	0.201 (4.39)	0.324 (3.19)
SH_{-1}	-0.048 (3.25)	-0.011 (1.47)	-0.116 (2.53)	0.006 (0.50)	-0.023 (3.41)	-0.329 (8.15)
X_{-1}	0.067 (4.13)	-0.019 (2.16)	0.03 (0.27)	0.026 (1.55)	0.018 (2.57)	-0.275 (4.11)
σ	0.039	0.019	0.12	0.045	0.021	0.078
Prob.persistence	0.912	0.945	0.920	0.939	0.983	0.847

See note to table 2.b.

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

Three results of our analysis with weekly averaged data deserve explanations: i) the long run anchor represented by the New York market before 1997; ii) the full disappearance of such an attractor over the subsequent period; and iii) the gradual emergence of the Hang Seng index as a new anchor since 1997.

During the pre-1997 regime only error correction with New York was at work. Such significant error correction must have been due to information flows. This regime was characterized by substantial volatility in returns on China's stock market (table 4) 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 millions of individual investors focused on the US stock market as a model and took into account what happened in that market in the long run. They must particularly have noticed that the domestic index was at a relatively low level (especially in the mid 1990s).

The vanishing influence of the New York anchor 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 (see figure A.1). 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-of-the-bubble phase was extremely rapid and concentrated over a short period. Trading volume in Shanghai 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 became 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 (State-Owned Enterprises) shares tradable.

Three complementary hypotheses can explain the switch to Hong Kong as a long run anchor in 1997. First the East Asian crisis may have represented the source of the change in regime. Such a crisis hit economic and financial relationships in East Asia in a deep and permanent way. Second, it may not be a pure coincidence that regime Two reasserted itself, after the interruption of the onset of the emerging market crisis (which started in the Czech republic in May 1997), right after the handover of Hong Kong to Mainland China, and almost exactly at the time of the (10%) collapse of the Hong Kong Stock market (23rd October 1997). However, these two reasons would not be able to explain why regime Two already emerged on the eve of the crisis, in early 1997. Such an emergence can only be explained by

the third factor, i.e. the development of multiple listing of Mainland firms on the Hong Kong stock market. The share of Hong Kong's market capitalisation by Mainland firms rose from 5% in 1992 to 16% in 1997, 21% in 1999 and 26% in 2002 (Xiao, 2005).

In the new millennium, the role played by the Hang Seng index 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. After mid-2001 the A-share market started to adjust downwards and the P/E ratios have been reduced from 50 or 60 to 20-30.

Since the middle of 2001 the A-share market in China has kept being a bear market. More and more market participants tried to connect this to a convergence of the Chinese stock market towards international stock markets, leading to a continuous fall in A-share prices. Investors paid more and more attention to 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 Hong Kong and domestic markets, investors started to worry that A-share prices and international prices would have to converge soon. As stressed by Chen and Kai (2004), in the past five years, Chinese investors have reduced their thought difference with international investors dramatically. The internationalisation of the investment strategy of Chinese investors will strongly support the convergence of share prices between Shanghai and Hong Kong. Such a process 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 in the prices of A-shares and H-shares will involve a fall in the former as well as a rise in the latter. Since 2003 the listing of Chinese Mainland companies in the Hong Kong market has intensified, but, due to the lack of transparency and credibility of such companies, international investors cannot be confident enough to buy their shares, explaining the low P/E ratios on the H-share market. However, Chinese Mainland investors know Chinese listed companies better than international investors. Capital flight is already used by Chinese residents to buy shares in Hong Kong,

including IPOs of Mainland firms listed in Hong Kong. When Chinese investors will be allowed to buy H-shares on a large scale, their price will rise.

VI. Conclusion.

Existing empirical work, using daily data up to 2002, has concluded that, in spite of high trade openness, China is not financially integrated, i.e. the A-share market is not linked, in the long run, with regional or international stock markets. Using either daily or mid-week closing prices and a single step Markov-switching error correction model, over a much longer sample (October 1992 through March 2005), for the Shanghai A-share market, we confirm such lack of cointegration with either the New York or Hong Kong market.

We build on some theoretical econometric work supporting the use of temporally aggregated data to test for the presence of long run relationships and the significance of error correction. The use of weekly averaged data over the same sample, with the same error correction model, provides evidence in favour of a long run relationship between the Shanghai A-share market index and the Standard and Poor's 500, up to late 1996. Subsequently a new (substitute) long run relationship gradually arose for the Shanghai market with the Hang Seng index. We provide explanations for such findings, on the basis of information flows, particularly in the case of the relationship with New York, but also in the case of Hong Kong, due to multiple listings of Mainland firms, as well as comparisons and gradual convergence of price-earnings ratios across the border. Since dynamic effects did not play a role at any time, our results echo those of Cheung et al. (2005) for the Chinese inter-bank market, similarly implying long run, but not short run, financial integration.

There are favourable prospects for the *de jure* 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 connections 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, since the P/E ratio of the same Mainland companies is still much larger in Shanghai than in Hong Kong.

Appendix I:
Stock price indices:
Shanghai-A, Hang Seng and S&P500

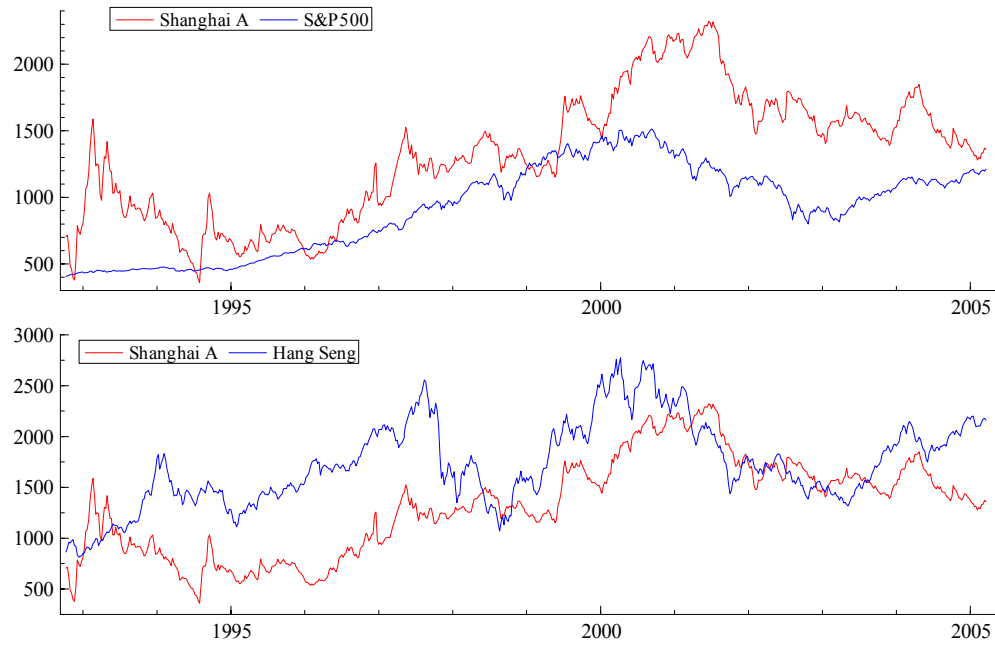


Figure A. 1: Weekly averaged indices for a) Shanghai-A and S&P500;
b) Shanghai-A and Hang Seng (adjusted range).

Appendix II:

Estimation of error correction models with daily data

Table A.2: Estimated coefficients of the Markov-switching error correction models:
Daily (closing price) data

Regime	with S&P			with Hang Seng		
	1	2	3	1	2	3
Intercept	-0.027 (2.00)	-0.011 (1.42)	0.200 (2.11)	-0.087 (2.91)	-0.021 (1.72)	0.324 (2.15)
ΔSH_{1t}	-0.038 (1.21)	0.016 (0.57)	-0.002 (0.04)	-0.04 (1.46)	0.017 (0.60)	-0.004 (0.06)
ΔSH_{2t}	-0.009 (0.29)	0.0003 (0.00)	0.044 (0.68)	-0.013 (0.43)	-0.000 (0.00)	0.04 (0.64)
ΔSH_{3t}	0.103 (3.17)	0.052 (2.07)	-0.009 (0.13)	0.100 (3.10)	0.051 (2.01)	-0.013 (0.19)
SH_{1t}	-0.004 (1.26)	-0.001 (0.71)	-0.04 (2.96)	0.0005 (0.24)	-0.0004 (0.31)	-0.035 (2.90)
X_{1t}	0.008 (2.63)	0.002 (1.56)	0.012 (0.93)	0.009 (2.46)	0.002 (1.51)	-0.008 (0.63)
σ	0.02	0.009	0.063	0.02	0.009	0.063
Probability of persistence	0.921	0.97	0.848	0.919	0.97	0.843

(t statistic); sample: 1992(October)- 2005(March), excluding public holidays in China when the Shanghai market is closed; SH= Shanghai-A index, X= Standard and Poor's 500 (columns 1 to 3) or Hang Seng index (col. 4 to 6).

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