One Date, One Break?*

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Abstract

This paper demonstrates that the conventional approach of using official liberalisation dates as the only existing breakdates could lead to inaccurate conclusions as to the effect of the underlying liberalisation policies. It also proposes an alternative paradigm for obtaining more robust estimates of volatility changes around official liberalisation dates and/or other important market events. By focusing on five East Asian emerging markets, all of which liberalised their financial markets in the late and by using recent advances in the econometrics of structural change it shows that (i) the detected breakdates in the volatility of stock market returns can be dramatically different to official liberalisation dates and (ii) the use of official liberalisation dates as breakdates can readily entail inaccurate inference. In contrast, the use of data driven techniques for the detection of multiple structural changes leads to a richer -and inevitably more accurate - pattern of volatility evolution emerges in comparison to focussing on official liberalisation dates.

Keywords: financial liberalisation, volatility, structural changes, breaks

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1. Introduction

The effects of financial liberalisation on stock market volatility have been the subject of controversy ever since emerging market economies began liberalising their financial markets in the 1980s and early 1990s.[1] Following Keynes[2], several authors have proposed that financial liberalisation could attract speculators and investors with short-term horizons, resulting in asset price bubbles and financial instability (e.g. Allen and Gale, 2000; Arestis and Demetriades, 1997, 1999; Singh, 1997, 2003). Stock market volatility could, however, decline following financial liberalisation if the number of traders increases (see Tauchen and Pitt, 1983). Empirical evidence on the subject is mixed, depending on the countries and periods that are studied, with recent studies showing that the outcome may depend on market characteristics, such as transparency and investor protection (e.g. Jayasuriya, 2005).

Previous empirical studies on the effects of financial liberalisation on stock market volatility implicitly assume that (i) there is a single break in the properties of the stock market returns; (ii) the timing of the break (breakdate) is known and coincides with the official stock market liberalisation date (or in some instances the announcement date)[3]. However, these assumptions are unlikely to be realistic for a number of reasons. With respect to (i), it is likely that there may be more than one break, which may or may not be directly due to stock market liberalisation. This could be because *financial liberalisation* is a broader concept than *stock market liberalisation*, in that it also includes other important areas of the financial system, such as the deregulation of banking activities, the lifting of interest rate controls, the removal of directed credit programmes, all of which have been widespread in emerging market economies.[4] It may also be due to changes in the political or

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[1] While the effects of financial liberalisation on stock market volatility are debatable, the view that increased volatility is undesirable is less controversial. Increased volatility is associated with higher capital costs and, consequently, lower investment; the latter may also decline because the 'option to wait' increases (Bekaert and Harvey, 1997). Empirical evidence provides some credence to this view. See, for example, Arestis *et al* (2001) who show that stock market volatility has negative effects on long run economic growth using quarterly data from five developed economies.

[2] Keynes (1964) regards liquidity as having destabilising effect on the market because of the assumption of market imperfection, particularly in relation to the availability of information to all participants.

[3] The only exception that we are aware of is the work of Cuñado et al. (2006) who use a data-driven technique to identify potentially existing breakdates.

[4] The typical sequencing of financial reforms in these economies usually starts from the lifting of interest rate controls and other banking restraints (see, for example, Arestis and

institutional environment, which may impact on investor behaviour. With respect to (ii), the breakdate in the data may or may not coincide with official liberalisation dates because financial market participants may adjust their behaviour well before or even after the official liberalisation dates, depending, for example, on the timing and credibility of announcements. For these reasons, the estimates of volatility changes due to financial liberalisation obtained by previous studies are likely to be biased or inefficient.[5]

In general, the purpose of this paper is twofold. First, to empirically demonstrate that the conventional approach of using official liberalisation dates as the only existing breakdates could lead to inaccurate conclusions as to the effect of the underlying liberalisation policies. Second, to propose an alternative paradigm for obtaining more robust estimates of volatility changes[6] around official liberalisation dates before attempting to isolate a causal relationship between them and the underlying financial reforms. To this end we propose the use of the *'Nominating-Awarding'* procedure of Karoglou (2010) for identifying the number and timing of structural breaks in volatility dynamics. Based on the identified breakdates we measure the unconditional volatility of each corresponding regime using both parametric and non-parametric volatility estimators and compare these values to those that are obtained by the conventional approach.[7] Our findings suggest that the conventional paradigm of analysing financial liberalisation policies based on pre and post liberalisation eras results in an 'averaging-out' of volatility patterns. This may well invalidate any inference about the effect that financial liberalisation policies and thus could prove misleading both to policy makers and to investors. Therefore, it is essential to robustify volatility estimates in the presence of multiple unknown breaks

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Demetriades, 1999). This may well result in breaks in stock market volatility since the shares of banks frequently represent a large fraction of stock market capitalisation.

[5] It is an established fact that not taking into account structural breaks in the estimation of GARCH-type models may result in over-estimating volatility persistence (Lamoureux and Lastrapes, 1990).

[6] Because we use three different estimators of volatility which are valid under different sets of assumptions, we prefer not to refer to these estimators as 'unbiased', even though at least one of these is likely to be depending on the true underlying data generating process, which is of course unknown.

[7] It is worth mentioning here that another approach to the one we adopt is a Markov Regime Switching specification. However, within the GARCH framework this raises considerable estimation concerns (see for example Bauwens, Preminger and Rombouts, 2006) which are naturally compounded both for many different regimes and for regimes with few observations such as transitional periods. Besides, it is inherently less flexible than the procedure proposed in this paper (for example in adopting a break-point that is suggested by exogenous information).

than simply rely on a conditional heteroskedastic structure such as the well-known (G)ARCH to capture the main components of volatility dynamics.

A key point of our paper is the selection of the emerging economies for our analysis, namely (South) Korea, Malaysia, Philippines, Taiwan and Thailand. These five East Asian countries liberalised their financial markets in the late 1980s or early 1990s and have been extensively studied in the broader literature on financial development, not least because of their importance to the world economy and the availability of reliable data.[8] They can therefore provide an excellent platform from which to highlight the importance of correctly identifying the number and timing of structural breaks prior to isolating a causal relationship between financial liberalisation reforms and volatility changes.

The paper is organised as follows. Section 2 outlines the statistical procedure for obtaining robust estimates of stock market volatility when multiple breaks may be present. Section 3 describes the data and data sources and provides the official financial liberalisation dates in each of the five East Asian countries studied, drawing on relevant literature. Section 4 presents the findings of the empirical application while Section 5 summarises and concludes.

2. Obtaining Reliable Volatility Estimates

An implicit, albeit easily recognized, assumption of the conventional paradigm is that a conditionally heteroskedastic model, typically the basic GARCH(1,1) of Bollerslev (1986), is all that is needed to capture the main components of volatility dynamics in the pre and post financial liberalisation periods. Since the main endeavour is the derivation of the unconditional variance as a proxy for volatility, the presence of breaks in volatility dynamics is not acknowledged as an important issue – breaks will have an impact on volatility persistence, not on the 'average' value of the unconditional variance. Consequently, having a sufficient number of observations for the pre and post liberalisation periods in order to get reliable estimates for the GARCH model is the main concern.

However, in order to obtain reliable volatility estimates around a financial liberalisation date it is essential to identify first potentially existing breaks in volatility dynamics before measuring volatility with various methods. The main reason is that

[8] See, for example, Demetriades and Luintel (2001) and Demetriades, Devereux and Luintel (1998).

an 'average' value of the unconditional variance in the pre and post liberalisation periods is very likely to ignore the impact of other, often temporary and country-specific, important market events. Consequently, in order to avoid drawing inference based on a distorted picture of volatility changes, it is necessary to avoid 'averaging-out' the impact of at least the most significant breaks in volatility dynamics.

The remainder of this section describes a method for obtaining robust to the presence of breaks volatility estimates. In particular, Sections 2.1 and 2.2 outline the '*Nominating-Awarding*' procedure of Karoglou (2010) for identifying the number and timing of all the potential breaks in each series while Section 2.3 explains what volatility measures can be used to address the small samples that are the result of segmenting the series. This last step becomes particularly important when conditioning on breaks due to the fact that even the basic GARCH(1,1) model yields significantly biased estimates for small samples (Hwang and Pereira, 2006), which in our case can be any identified segment with few observations. Consequently, the use of alternative measure of volatility in such cases becomes imperative.

2.1. The 'Nominating breakdates' stage

The 'Nominating breakdates' stage means that a certain procedure based on one or more statistical tests identify some dates as possible breakdates. In recent years, a number of statistical tests have developed for that reason with considerably different properties (see for example, Sansó et al., 2003). In this paper, we use a battery of these tests in order to take advantage of the special characteristics of each of them and particularly the trade-off between size distortions and low power. In fact, discrepancies in the detected breakdates can be quite informative of the properties of the underlying stochastic process. Interestingly, however, the detected breakdates are surprisingly similar. For the purposes of this paper, we use the following tests:

(a) I&T (Inclàn and Tiao, 1994)
(b) SAC_1 (The first test of Sansó, Aragó, and Carrion, 2003)
(c) SAC_2^{BT}, SAC_2^{QS} (The second test of Sansó, Aragó, and Carrion, 2003, with the Bartlett kernel and the Quadratic Spectral kernel correspondingly)[9]
(d) K&L (The refined by Andreou and Ghysels, 2002 version of the Kokoszka and Leipus, 2000 test with the Vector Autoregressive Heteroskedasticity and Autocorrelation Consistent or VARHAC kernel of den Haan and Levin, 1998).

[9] The Newey and West (1987) automatic procedure provides the bandwidth selection for the Bartlett and Quadratic spectral kernels.

There are a number of reasons why these tests have been selected to identify the structural changes in each of the series. First, although all of these tests are designed to detect a structural change in the volatility dynamics, Karoglou (2006) shows that many CUSUM-type tests (including all the above) do not discriminate between shifts in the mean and shifts in the variance. This is a plausible feature since all types of breaks need to be considered in order to determine if and to what extent the distributional properties change when moving from one regime to another. A second reason for selecting these CUSUM-type tests is that their properties for strongly dependent series have been extensively investigated (e.g. Andreou and Ghysels, 2002, Sansó, Aragó, and Carrion, 2003, Karoglou, 2006) and there is evidence that they perform satisfactorily under the most common ARCH-type processes. Thus, even when there is a break in a conditionally heteroskedastic process these tests can detect it. That is, the tests do not exhibit size distortions and they have considerable power even when the assumption of within-segment homoskedasticity is relaxed in order to include ARCH-type structures. In fact, (c) and (d) have some plausible properties even in the presence of IGARCH effects. Nevertheless, Karoglou (2006) shows that the relative performance of each of the above tests depends on the underlying data generating process (DGP). Consequently, since the true DGP is not known, it is preferable to use them all and select the breakdate according to an appropriate set of rules.

The above tests can also be used to identify multiple breaks in a series. This is achieved by incorporating the breaks in an iterative scheme (algorithm) and applying them to sub-samples of the series. The employed algorithm comprises of the following six steps:

1. *Calculate the test statistic under consideration using the available data.*
2. *If the statistic is above the critical value split the particular sample into two parts at the date at which the value of a test statistic is maximized.*
3. *Repeat steps 1 and 2 for the first segment until no more (earlier) change-points are found.*
4. *Mark this point as an estimated change-point of the whole series.*
5. *Remove the observations that precede this point (i.e. those that constitute the first segment).*
6. *Consider the remaining observations as the new sample and repeat steps 1 to 5 until no more change-points are found.*

The algorithm is implemented with each of the (single breakdate CUSUM-type) test statistics described above (i.e. I&T, SAC_1, SAC_2^{BT}, SAC_2^{QS}, K&L) and is applied to each series in ascending and descending time order so as to avoid potentially existing masking effects. The main feature of the algorithm that differentiates it from a simple binary division procedure, such as the ICSS procedure, is that it ensures that the existing breaks are detected in a time-orderly fashion. In other words, the first break proposed by the algorithm is also the earliest break in the series, the second break proposed is the second earliest break, and so forth. This is particularly important when transitional periods exist in which case a simple binary division procedure is likely to produce more breaks in the interim period. In the absence of transitional periods both procedures will produce the same breaks. It is worth noting that the breakdates obtained when the procedure was applied in ascending and descending time order are almost identical which is an additional piece of evidence regarding the robustness of the procedure. The nominated breakdates for each series are simply all those which have been detected in each case unless (i) they are detected only by the I&T since this tests suffers from considerable size distortions for non-mesokurtic data such as the stock market data (ii) the resulting segments will have less than 50 observations – a rule which has been applied in only one instance at the end of the sample. Note that at this stage we are not much concerned with detecting more breaks than those that actually exist because whichever is not an actual breakdate will be picked up in the 'Awarding breakdates' stage.

2.2. The 'Awarding breakdates' stage

Following the nomination of the potential breakdates outlined in the preceding section, this stage serves as a screening process of the detected breakdates. The 'Awarding breakdates' stage describes the procedure that is used to decide whether a nominated breakdate is indeed a breakdate. Initially, we assume that all nominations are breakdates. In fact, one can infer that this is what is being done in the empirical literature so far. We then unite contiguous nominated segments (i.e. segments that are defined by the nominated breakdates) unless the variances of the contiguous segments are statistically different (as suggested by the battery of tests which is described below). This testing procedure is repeated until no more segments can be united.

Overall, the battery of tests involves several methods designed to test for the homogeneity of variances of different samples, which in this case are two contiguous segments. These tests constitute a different approach to the tests described in the

‘Nominating breakdates’ stage in that they test for the homogeneity of variances without encompassing the time-series dimension of the data. They include the standard F-test, the Siegel-Tukey test with continuity correction (Siegel and Tukey, 1960, and Sheskin, 1997), the adjusted Bartlett test (see Sokal and Rohlf, 1995, and Judge, et al., 1985) and the Levene test (1960).

In short, the F-test requires equal sample sizes and is sensitive to departures from normality. On the other hand, the Siegel-Tukey test is based on the assumption that the samples are independent and have the same median. The Bartlett test is also robust when the sample sizes are not equal despite it still being sensitive to departures from normality. Its adjusted version makes use of a correction factor for the critical values and the arcsine-square root transformation of the data to conform to the normality assumption. Finally, the Levene test is an alternative to the Bartlett test albeit less sensitive to departures from normality.

2.3. Volatility Estimators

The last step of the method involves measuring volatility and specifically the magnitude and direction of the change in volatility. Following the empirical literature, we proxy volatility in each segment by the underlying unconditional variance. The conventional paradigm is mainly focused on a parametric approach and specifically the GARCH approach to obtain volatility estimates. However, Hwang and Pereira (2006) have shown that such an approach can provide reliable volatility estimates only when there are more than 500 observations available. When breaks are taken into account, the subsequent segmentation of the sample space is most likely to result in segments with fewer observations.[10] In such cases, a GARCH estimate is likely to be biased. To address this issue, we propose the use of Heteroskedasticity and Autocorrelation Consistent (HAC) estimators of the variance and particularly the use of the Vector Autoregressive or VARHAC kernel of den Haan and Levin, 1998 which has the additional advantage that circumvents the problem of selecting an appropriate bandwidth unlike other kernels. Nevertheless, for comparison purposes, we also

[10] For this reason we report volatility estimates from the simpler GARCH structure rather than the APARCH-in-mean structure which we have used in an earlier draft of this paper. Nevertheless, the results are almost identical which to a large extent is attributed to the fact that in the identified segments, asymmetric and in-mean effects were statistically insignificant.

calculate the sample standard deviation as well as the square root of the unconditional variance of the best fitting GARCH whenever this is estimable.[11]

3. East Asian Data and Liberalisation Dates

It is widely accepted that the conditional mean of the returns exhibits little predictability from the past (Bekaert and Harvey, 1997). However, we also consider the possibility of moving average error terms induced by calendar effects. We therefore follow the procedure suggested by Pagan and Schwert (1990) to remove potential day-of-the-week effects.

The data used in this paper are the continuously compounded daily stock returns obtained by the daily closing stock price indexes, expressed in the local currency[12] of: (i) Korea Stock Price Index; (ii) Taiwan Weighted Stock Index; (iii) Kuala Lumpur Composite Index; (iv) Stock Exchange of Thailand Index and (v) the Philippines Stock Exchange Composite Index. The sample period spans four years before and after official financial liberalisation dates. The data is obtained from *Datastream*.

Financial Liberalisation Dates of East Asian Emerging Markets

The selection of the official liberalisation dates for five Asian emerging markets draws on the following papers: Santis and Imrohoroglu (1997), Henry (2000), Kim and Singal (2000), Bekaert and Harvey (2000), Fuchs-Schundeln and Funke (2001), Kassimatis (2002) and Bhattacharya and Daouk (2002). Table 1 summarises the liberalisation dates used in each of these papers and lists those that we adopt in the rest of this paper, namely the ones that most papers agree on. These are as follows: January 1992 for South Korea, January 1991 for Taiwan, December 1988 for Malaysia, September 1987 for Thailand and June 1991 for the Philippines.

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[11]We define the 'best fitting GARCH' as the GARCH with the largest log-likelihood and with statistically significant coefficients (at 5% significance level). Information criteria are not used mainly because of their limited ability to identify the true structure of GARCH-type processes (Mitchell and McKenzie, 2003).
[12] US dollar indexes are not employed in order to avoid introducing exchange rate volatility effects.

Table 1: Official Financial Liberalisation Dates in East Asia

Source / Country	*Santis and Imrohoroglu (1997)*	*Henry (2000)*	*Kim and Singal (2000)*	*Bekaert and Harvey (2000)*
Korea	Jan-92	Jun-87	Jan-92	Jan-92
Malaysia	Dec-88	May-87	Dec-88	Dec-88
Philippines	Oct-89	May-86	Jul-86	Jun-91
Taiwan	Jan-91	May-86	Jan-91	Jan-91
Thailand	Dec-88	Jan-88	Aug-88	Sep-87
Work of / Country	*Fuchs-Schundeln and Funke (2001)*	*Kassimatis (2002)*	*Bhattacharya and Daouk (2002)*	***ADOPTED***
Korea	Jan-92	Jan-92	Jan-92	***Jan-92***
Malaysia	Dec-88	NA	Dec-88	***Dec-88***
Philippines	Jun-91	Nov-91	Jun-91	***Jun-91***
Taiwan	Jan-91	Jan-91	Jan-91	***Jan-91***
Thailand	Sep-87	NA	Sep-87	***Sep-87***

Table 2 presents some descriptive statistics for the stock returns in these markets for the full sample and the two sub-sample periods defined by the official liberalisation dates. Based on this preliminary description of the data, it appears that (i) the mean of stock returns increased in the cases of Korea, Malaysia and the Philippines following financial liberalisation while it decreased in the cases of Taiwan and Thailand; (ii) the stock return volatility (as measured by the standard deviation) appears to have declined after liberalisation in four of the five countries, the exception being Thailand, where it appears to have increased considerably.

Table 2: Descriptive Statistics of Stock Returns

		Period	**Mean**	**St. Deviation**	**Skewness**	**Kurtosis**	**Observations**
Korea	*Full Sample*	(Jan 88 – Dec 95)	0.0108	0.5969	0.2929	2.9882	2086
	Pre-Lib	(Jan 88 – Dec 91)	0.0063	0.6231	0.1928	3.0126	1043
	Post-Lib	(Jan 92 – Dec 95)	0.0153	0.5694	0.4275	2.8626	1043
Malaysia	*Full Sample*	(Dec 84 – Nov 92)	0.0158	0.6326	-2.0698	24.6773	2086
	Pre-Lib	(Dec 84 – Nov 88)	0.0062	0.7442	-2.1061	21.6072	1043
	Post-Lib	(Dec 88 – Nov 92)	0.0254	0.4967	-1.4075	18.6465	1043
Philippines	*Full Sample*	(Jun 87 – May 95)	0.0314	0.8577	-0.1761	10.5878	2088
	Pre-Lib	(Jun 87 – May 91)	0.027	1.0506	-0.2043	8.5135	1045
	Post-Lib	(Jun 91 – May 95)	0.0357	0.6064	0.1088	2.2532	1043
Taiwan	*Full Sample*	(Jan 87 – Dec 94)	0.04	1.04	-0.0905	1.8429	2087
	Pre-Lib	(Jan 87 – Dec 90)	0.0613	1.2187	-0.1363	0.9397	1043
	Post-Lib	(Jan 91 – Dec 94)	0.0187	0.8237	-0.0161	2.8274	1044
Thailand	*Full Sample*	(Sept 83 – Aug 91)	0.0329	0.6073	-0.8298	11.5289	2087
	Pre-Lib	(Sept 83 – Aug 87)	0.0371	0.2689	0.1676	8.1141	1043
	Post-Lib	(Sept 87 – Aug 91)	0.0287	0.8157	-0.6779	5.8025	1044

4. Empirical Application

This penultimate part provides the empirical evidence that underscores the weaknesses of the conventional paradigm in measuring the impact of financial liberalisation reforms and emphasizes the importance of drawing inference only after conditioning on multiple unknown breaks. Section 4.1 contains the results from applying the 'Nominating breakdates' stage procedure; Section 4.2 contains the results from the 'Awarding breakdates' stage procedure; and Section 4.3 briefly compares the different volatility estimates. Finally, Section 4.4 discusses analytically for each country the evolution of volatility.

4.1. The 'Nominated' breakdates

Table 3 reports the results of applying the procedure described in the 'Nominating breakdates' stage in Section 2. Overall we identify three to four breaks in each series which results to four or five segments correspondingly. With the exception of only one detected breakdate, all other detected breakdates define segments that involve more than 50 observations but often with less than substantially less than 200 observations.

Table 3: Detected Structural Changes

	datapoint	I&T	SAC_1	SAC_2^{BT}	SAC_2^{QS}	K&L	adopted
korea	597	√	√	√*	√	√	yes (16-04-90)
	1291	√	√	√*	√	√	yes (10-12-92)
	1607	√	√*	-	√*	√*	yes (01-03-94)
	1828	√*	-	-	-	-	no
	1873	√*	-	-	-	-	no
malaysia	751	√	√	-	√	√*	yes (19-10-87)
	818	√	√	-	√	√*	yes (19-01-88)
	1756	√	-	-	-	-	yes (26-08-91)
	1818	√*	-	-	-	-	no
phillipines	110	√	-	-	-	-	no
	149	-	√	√	√	√	yes (20-12-87)
	1128	√	√	√	√	√	yes (25-09-91)
	1656	√	√	√*	√	√*	yes (04-10-93)
	1810	√	√	√*	√	√*	yes (06-05-94)
	1952	√	-	-	-	-	no
	2037	√*	-	-	-	-	no
taiwan	848	√	√	√	√	√	yes (02-04-90)
	1094	√	√	√	√	√	yes (12-03-91)
	1259	√	√	√*	√	√*	yes (29-10-91)
	1558	√	-	-	-	-	no
	1647	√	-	-	-	-	no
	1803	√	-	-	-	-	no
	1875	√	-	-	-	-	no
	2025	√	-	-	-	-	no
	2046	√	-	-	-	-	no
thailand	781	√	√	√	√	√	yes (28-08-86)
	1805	√	√	√	√	√	yes (01-08-90)
	1955	√	√	√*	√	√	yes (27-02-91)
	2044	√	√	-	√	-	no

Note: √ denotes statistical significance at 1% level, √* at 5% level, and - no statistical significance.

4.2. The ‘Awarded’ breakdates

Table 4 reports the results of carrying out the procedure described in the ‘Awarding breakdates’ stage in Section 2. The same table also reports the results of applying the same procedure to the segments defined by the official liberalisation dates. The results confirm that the neighbouring segments as determined by the adopted breakdates have statistically different variances. The same tests also suggest that the pre and post liberalisation periods do not always yield statistically different variances at the 1% demonstrating vividly the effect of ‘averaging-out’ volatility changes – in the case of Korea three of the tests suggest no variance change after the official liberalisation date.

Table 4: Robustness Tests

		F-statistic	Siegel-Tukey	Bartlett	Levene	change in variance
korea	before & after liberalisation	1.19†	0.41†	7.69	0.95†	no
	Regime 1 & 2	2.21	4.94	96.03	38.85	yes
	Regime 2 & 3	1.93	2.56	42.94	15.02	yes
	Regime 3 & 4	1.50	2.58	15.92	10.70	yes
malaysia	before & after liberalisation	2.21	8.60	159.66	48.42	yes
	Regime 1 & 2	10.21	4.89	296.10	114.05	yes
	Regime 2 & 3	13.38	6.96	420.20	166.18	yes
	Regime 3 & 4	2.41	4.60	52.77	37.38	yes
phillipines	before & after liberalisation	3.02	5.97	303.13	68.56	yes
	Regime 1 & 2	6.75	8.61	344.44	176.72	yes
	Regime 2 & 3	2.27	5.50	103.82	42.33	yes
	Regime 3 & 4	2.41	4.60	52.77	37.38	yes
	Regime 4 & 5	2.13	3.15	29.69	18.67	yes
taiwan	before & after liberalisation	2.13	10.61	145.00	114.14	yes
	Regime 1 & 2	3.69	10.67	195.62	201.98	yes
	Regime 2 & 3	3.45	7.27	65.98	66.59	yes
	Regime 3 & 4	2.06	4.86	41.89	32.13	yes
thailand	before & after liberalisation	9.09	16.44	1072.55	265.60	yes
	Regime 1 & 2	9.51	16.13	873.02	224.07	yes
	Regime 2 & 3	5.55	8.31	278.93	170.04	yes
	Regime 3 & 4	4.11	4.53	62.88	30.82	yes

Note: † denotes the absence of statistical significance at 10% level. In all other cases, the statistical significance is found below 1% level.

4.3. Volatility Estimates[13]

Table 5a and 5b depict for each economy the three volatility estimates for both the pre- and post-liberalisation periods and the segments that are defined by the identified breaks.

[13] A series of figures that illustrate the results of this and of the next section are available from the webpage of the corresponding author.

Table 5a: volatility estimates when breaks are taken and not taken into account

	From	To	Obs.	Std. Dev.	VARHAC	GARCH		From	To	Obs.	Std. Dev.	VARHAC	GARCH
Korea	01/01/88	31/12/91	1459	62.4%	59.7%	62.2%	Malaysia	03/12/84	30/11/88	1459	73.6%	80.8%	68.7%
	01/01/92	29/12/95	1459	57.2%	59.5%	56.1%		01/12/88	30/11/92	1459	49.5%	48.2%	49.2%
	01/01/88	15/04/90	596	51.1%	51.1%	51.2%		03/12/84	18/10/87	750	60.5%	60.6%	61.8%
	16/04/90	09/12/92	693	76.0%	68.1%	77.6%		19/10/87	18/01/88	66	193.4%	138.2%	94.6%
	10/12/92	28/02/94	318	54.7%	54.5%	54.7%		19/01/88	25/08/91	939	52.9%	52.8%	51.9%
	01/03/94	29/12/95	479	44.6%	44.7%	44.6%		26/08/91	30/11/92	331	33.6%	31.8%	33.8%
Taiwan	01/01/87	31/12/90	1459	120.2%	128.4%	131.4%	Thailand	01/09/83	31/08/87	1459	26.4%	35.5%	66.1%
	01/01/91	31/12/94	1459	82.4%	73.5%	75.6%		01/09/87	31/08/91	1459	79.6%	85.5%	inf.
	01/01/87	01/04/90	1093	122.8%	125.5%	137.2%		09/01/83	27/08/86	780	19.1%	20.1%	21.2%
	02/04/90	11/03/91	246	185.9%	186.2%	185.7%		28/08/86	31/07/90	1024	58.8%	63.4%	72.1%
	12/03/91	28/10/91	165	100.0%	100.0%	107.2%		01/08/90	26/02/91	150	138.4%	134.2%	138.3%
	29/10/91	30/12/94	829	69.6%	60.7%	69.7%		27/02/91	30/08/91	133	68.3%	58.1%	63.6%

Table 5b: volatility estimates when breaks are taken and not taken into account

	From	To	Obs.	Std. Dev.	VARHAC	GARCH
Philippines	01/06/87	31/05/91	1459	103.8%	103.9%	93.9%
	01/06/91	31/05/95	1459	59.7%	67.1%	61.1%
	01/06/87	19/12/87	145	201.4%	178.1%	206.1%
	20/12/87	24/09/91	982	77.6%	83.1%	77.0%
	25/09/91	03/10/93	528	51.5%	51.6%	54.3%
	04/10/93	05/05/94	154	80.0%	101.9%	80.1%
	06/05/94	31/05/95	279	54.8%	54.8%	54.8%

Note: the shaded entries involve the pre- and post-liberalisation periods as defined by the official financial liberalization date.

In general, for each country, we obtain results using the three alternative measures of volatility in (i) the pre and post liberalisation periods and (ii) each of the identified regimes. Overall, the three volatility estimates in the first case (i) are quite similar except from Thailand which yields dramatically different values both in the pre and post liberalisation periods. When instead we look at the volatility estimates in the second case (ii) we observe that by conditioning on multiple unknown breaks the three volatility estimates are substantially different mainly for the smaller segments. Finally, if we contrast the values of each volatility estimator when breaks are not taken into account to the corresponding values when breaks are taken into account the VARHAC volatility estimator discussed in Section 2 ‘averaged-out’ the impact of breaks substantially better than the other two estimators. In other words, the volatility estimates that it provides when breaks are not taken into account deviate the least from the values that it provides when breaks are taken into account – and which are similar to the values of the other estimators. Therefore the theoretical advantages of the VARHAC volatility estimator appear to be of practical importance.

At this point it is worth mentioning that given the fact that we have already detected breaks in the Thailand series, there are substantial differences between the three volatility estimates when no breaks are taken into account, and this can be viewed as a sign of breaks in volatility dynamics that are not 'averaged-out' effectively. In other words, the fact that the three volatility estimates in one series in the pre- and post-liberalisation periods yield substantially different results, while for large samples the three volatility estimators are expected to yield identical values, suggests that the samples in each of these periods are too small to manifest this asymptotic property. Consequently, a key element of analysis of the traditional paradigm (to make use of about 1000 observations for each period) is empirically shown to be sample-specific and thus unreliable.

It is also worth mentioning that when we look at the three volatility estimates in the post liberalisation period of the Thailand series, it becomes apparent that the inflated persistence of the estimated GARCH model due to the presence of breaks that are not taken into account has even led to the manifestation of Integrated GARCH effects. In other words, and in accordance to the findings of Lamoureux and Lastrapes (1990), within the GARCH framework, the post-liberalisation period when breaks are not taken into account is a sequence of realisations of a process with infinite volatility despite the fact that this is not actually the case. When compared to the corresponding values when breaks are taken into account, it becomes evident that the VARHAC volatility estimate has actually 'averaged-out' the impact of the unaccounted for breaks substantially more accurately.

4.4. The evolution of volatility

When we look at the volatility measures in Korea we see that the estimated measures of volatility before and after the official liberalisation date of 1 January 1992 have declined slightly. The GARCH-derived estimate shows a decline of 9.8%, the standard deviation a decline of 8.3% and the VARHAC estimate shows a marginal decline of 0.3%.[14]

In contrast, when we look at the same volatility measures when breaks are taken into account a much richer evolution of volatility in the pre and post liberalisation periods is revealed. The volatility measures in the first segment, which covers the period 1 January 1988 – 15 April 1990, was, in fact considerably lower than suggested by the results when breaks are not taken into account. In the second

14 Note, however, that the tests reported in Table 4 suggest that these changes are not statistically significant in this particular case.

segment, which covers a twenty month period before the official liberalisation data and an eleven month period after the official liberalisation date, volatility increased substantially: the GARCH measure shows an increase of 51.6%, the standard deviation an increase of 48.7% and the VARHAC an increase of 33.3%. The third segment, however, which starts almost a year after the official liberalisation date is one of decreasing volatility, with the three measures decreasing by 29.5%, 28% and 20% respectively. Finally, the fourth segment which starts twenty six months after the liberalisation date exhibits a further decline in volatility of 18.5% in both the first two measures and 18.0% in the third. As a result, a comparison of the first and fourth segment shows that volatility has declined by around 12.7% (12.9%, 12.7% and 12.5%, respectively).

A plausible interpretation of the Korean results suggests that the first regime corresponds to the period before any news regarding stock market liberalisation had reached the market while the second regime corresponds to the time when information about liberalisation reached market participants, creating uncertainty. It is interesting, however, that the second regime continues well after the official liberalisation date. Even in the third regime, which begins eleven months after the liberalisation date, uncertainty appears to be higher than in the first regime. It takes more than two years after the official liberalisation date before uncertainty is reduced to pre-liberalisation levels. Thus, focusing on the regimes that are based on the official liberalisation dates completely masks this rich volatility pattern.[15]

A similar conclusion, if more pronounced, emerges when analysing the results for Malaysia. When breaks are not taken into account it appears that liberalisation led to a decline in volatility of between 28.4% and 40.3%, depending on which measure is used. However, when breaks are taken into account a much more striking evolution of volatility is revealed. Volatility increases very substantially for a period of three months, about a year before the official liberalisation date. The standard deviation suggests an increase in volatility of 219.7% while VARHAC shows an increase of 128.1% and GARCH a smaller increase of 53.1%, which is nevertheless also rather large. About a year before the liberalisation date of 1 December 1988

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[15] It may also be argued that the four regimes found for Korea using data driven techniques correspond to different *financial* as opposed to *stock market* liberalisation periods. The financial liberalisation indices constructed by Abiad and Mody (2005), which take on board credit controls, interest rate controls, entry barriers, regulations and privatization, as well as controls on international transactions, would however suggest only three different regimes for Korea: 19986-88, 1989-90, and 1991-96.

volatility declines quite substantially and remains low for a period of three and a half years: the GARCH measure shows a decline of 45.1%, the standard deviation a decline of 72.6% and the VARHAC a decrease of 61.8%. A further decline in volatility, in the range of 35-40% depending on the measure used, occurs in the fourth regime, which starts approximately two years and nine months after the official liberalisation date. As a result, volatility exhibits a decline in the range of 45.3-47.5%, depending on measure used, when the first and the last (fourth) regimes are compared.

The Philippines exhibits an even richer evolution of volatility, given that there are five different regimes. When breaks are not taken into account there appears to be a decline in volatility in the post liberalisation period that ranges from 34.9% in the case of the GARCH measure to 42.5% in the case of the standard deviation. However, when breaks are taken into account it becomes obvious that this masks a much more considerable drop in volatility when one compares the first regime with the last (fifth) one, which ranges between 69.2% and 73.4% depending on the measure used. In between the first and fifth regimes there are two consecutive periods of declining volatility, followed by a period of increasing volatility, ending with a period of declining volatility. The official liberalisation date falls three months before the end of the second regime. The period of increased volatility, which lasts for about seven months, occurs more than two years after the official liberalisation date.

The case of Taiwan is very similar to that of Malaysia and to some extent, Korea. The pre-liberalisation period includes a regime of substantially increased volatility which starts about nine months before the official liberalisation date and ends three months after. The increase in volatility ranges from 35.3% in the case of the GARCH measure to 51.4% for the standard deviation. This period is then followed by two regimes of declining volatility, lasting about seven months and more than three years, respectively. The decline in volatility between the first and fourth regimes ranges from 43.3% to 51.6% depending on which measure is used. Comparing the pre and post liberalisation periods shows a decline in volatility in the range of 31.4% to 42.8%, which masks all the aforementioned changes.

Thailand presents a sharp contrast to the other countries in that the results suggest an increase in volatility, following the financial liberalisation of 1 September 1987. The comparison of the pre and post liberalisation periods shows an increase of 201.5% for the standard deviation and 140.8% for the VARHAC measure. The GARCH measure indicates a change to an infinite unconditional variance, which further illustrates the limitations of artificially imposing a single breakdate in the sample period. All volatility measures suggest how that volatility more than trebled about a year before the official liberalisation date. This regime continues for almost three years after the liberalisation date. Moreover, it is followed by a seven-month period where volatility increases by 91.8%-135.4%, depending on the measure used. In the final period, which lasts about six months, volatility declines by about 50%, but this is not sufficient to bring it back to its pre-liberalisation level. In fact, comparison of the first and last regimes suggests that volatility increased by 189.1%-257.6%, depending on the measure employed. Once again, a before and after comparison masks several important volatility swings.

5. Conclusions

This paper highlights the importance of correctly identifying the number and timing of structural breaks when analysing changes in stock market volatility that may be directly or indirectly related to financial liberalisation. The volatility dynamics that emerge when breakdates are carefully extracted from the data are much richer than those suggested by the official liberalisation dates. In three of the five countries analysed - Korea, Malaysia and Taiwan – volatility increases before the official liberalisation date and subsequently declines below its original level. Focussing only on the pre and post stock market liberalisation period altogether fails to detect a period of increased volatility, which in the case of Korea exceeds two years. In the case of the Philippines, analysing the pre and post liberalisation periods, masks an initial marked decline in volatility and fails to pick up a period of substantially increased volatility that occurs more than two years after the official liberalisation date. In the case of Thailand, focussing on the official liberalisation date fails to pick up a decline in volatility that occurs in the fourth (final) regime, which nevertheless is not sufficient to reduce volatility to its pre-liberalisation level. In all cases the analysis of pre and post liberalisation periods results in an 'averaging-out' of volatility patterns. Thus, important changes in volatility are not be detected, resulting in inaccurate inference, potentially misleading policy makers and investors.

In the case of Korea for example, the conventional approach to analysing volatility around a financial liberalisation reform would lead policy makers and investors to believe that the 1992 financial reform had no impact on the riskiness of the Korean stock market suggesting that the market has maintained its main (statistical) characteristics. However, the approach we propose in this paper reveals two substantial volatility swings which indicate that the 1992 financial reform, took place officially at about the middle of a two and a half years period characterised by an almost fifty percent increase in volatility. After that period the volatility of the Korean stock market dropped about 13% from its starting levels. Consequently, if one accepts that the detected breaks are causally linked to the opening of the Korean stock market to foreign investors, then our approach would suggest that the main (statistical) characteristics of the Korean stock market have been significantly altered by the 1992 financial reform. If we do not make such a (rather strong) assumption then our approach would suggest that an analysis of the reasons behind the entry into each new volatility regime is paramount in correctly assessing the impact of the 1992 financial reform. Therefore, the conventional approach would mask the true benefits/costs of the reform in terms of stock market volatility to policy makers; it would lead investors who are interested in encompassing this stock market in their international diversification strategy to miscalculate the actual impact of such an action to the performance of their portfolios. Similar arguments can be made for the other economies we studied. In conclusion, our findings suggest that the analysis of the effects of financial liberalisation on stock market uncertainty remains fertile ground for further research.

An area where further research would be fruitful would be to examine whether the different volatility regimes that we detect using data driven techniques correspond to - or are indeed caused by - broader financial reforms, which are not directly linked to the opening of stock markets to foreign investors, such as the relaxation of credit and interest rate controls, entry barriers in banking, financial sector privatisation etc. The financial liberalisation literature now provides indices of financial liberalisation, albeit at frequencies that do not match well with the frequency of stock market returns.[16] Such an exercise would therefore require some additional data work as well as novel econometric approaches to address the causality issue.

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[16] See for example Abiad and Mody (2005), who provide such an index for 35 economies, on an annual basis for the period 1973-1996.

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