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FOUR ESSAYS ON MONETARY AND FINANCIAL INTEGRATION IN ASIA

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présentée par

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Résumé

Depuis la crise financière asiatique de 1997-98, plusieurs propositions visant à renforcer l'intégration monétaire des pays asiatiques ont vu le jour, comme l'adoption d'un ancrage sur un panier de devises, le retour à un ancrage souple sur le dollar, la création d'une unité monétaire asiatique (ACU) ou l'adoption d'une monnaie unique. Dans cette thèse, nous proposons quatre contributions originales à l'étude de l'intégration monétaire et financière des pays asiatiques.

Dans le premier chapitre nous déterminons la sensibilité relative des devises asiatiques (ASEAN-5, Corée du Sud) face aux chocs simulés sur le dollar, l'euro et l'ACU. Nous mettons en évidence la volonté de ces pays de se détourner d'une politique de change exclusivement centrée sur le dollar vers une politique plus flexible, où le poids de l'ACU semble avoir gagné en importance.

Le deuxième chapitre met l'accent sur la synchronisation entre les cycles des affaires de l'ASEAN-5. Nous montrons que la corrélation entre les cycles est plus forte durant les phases de contraction mais que la dynamique d'ajustement est propre à chaque pays. Par ailleurs, certains cycles des affaires de l'ASEAN-5 contiennent des informations pertinentes pour prédire les changements de régime des autres pays.

Le troisième chapitre examine le co-mouvement entre les taux de change réels de l'ASEAN-5 du point de vue de la parité de pouvoir d'achat généralisé (Enders and Hurns , 1994, 1997). Nous montrons que les taux de change réels sont liés par un processus à mémoire longue, ce qui soutient l'idée d'une intégration monétaire plus poussée entre différents sous-groupes de pays.

Enfin dans le dernier chapitre, nous examinons le degré d'intégration des marchés boursiers en Asie (ASEAN-5, Hong Kong, Japon). Nos résultats montrent que la volatilité des marchés boursiers internationaux partagent une tendance stochastique commune. En revanche, les marchés boursiers des pays émergents apparaissent encore segmentés tant au niveau global que régional.

Mots clés: intégration monétaire, intégration financière, zone monétaire optimale, Asie, unité monétaire asiatique, taux de change réel, cycle des affaires, volatilité des marchés boursiers, cointégration fractionnaire, modèle Markov-switching, probabilités de transition *time-varying*.
JEL: C32, E32, F33, F36, F41.

Abstract

Since the 1997-98 Asian financial crisis, several proposals have emerged to strengthen monetary integration in Asia, such as a *de facto* currency basket arrangement, the return to the soft dollar pegging, the creation of an Asian Currency Unit (ACU) or the adoption of a single currency. This thesis proposes four contributions to the study of Asian monetary and financial integration.

The first chapter examines to what extent the East Asian exchange rates (ASEAN-5, South Korea) are sensitive to shocks simulated on the US dollar, the euro and the ACU. We show that these countries have moved from a US dollar-based pegging system to a more flexible exchange rate policy, where the weight of the ACU has increased over the last years.

The second chapter attempts to analyze the correlation among the ASEAN-5 business cycles. Estimates reveal that correlations are higher during downturns but the process of adjustment to shocks displays idiosyncratic features. We also provide evidence that the signals contained in some leading ASEAN-5 business cycles, in contrast to global cycles (China, Japan and United-States), can impact the ASEAN-5 individual business cycles.

The third chapter examines the co-movement among the ASEAN-5 real exchange rates through the generalized purchasing power parity (Enders and Hurns , 1994, 1997). We find that real exchange rates are tied through a long memory process, supporting further monetary integration among different sub-groups of the ASEAN-5.

In the last chapter, we investigate to what extent the stock markets in Asia (Hong Kong, Japan, ASEAN-5) are integrated. Our results reveal that the stock market volatilities in developed countries share a common stochastic trend. Conversely, emerging markets appear to be segmented from both each other and global markets.

Keywords: monetary integration, financial integration, optimum currency area, ASEAN-5, Asian Currency Unit, real exchange rate, business cycle, stock market volatility, fractional cointegration, Markov-switching model, time-varying transition probabilities.
JEL: C32, E32, F33, F36, F41.

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Introduction Générale

L'objet de cette thèse est d'étudier l'intégration monétaire et financière des pays asiatiques afin de déterminer si une stabilisation coordonnée des devises asiatiques est envisageable dans un futur proche. Le concept d'intégration monétaire recouvre toute forme d'arrangements susceptible d'atteindre cette stabilisation, notamment, l'adoption d'une monnaie unique mais aussi un ancrage coordonné sur une même devise ou sur un panier de devises. Tout au long de cette étude, nous utilisons plusieurs outils empiriques appropriés, en vue de dépasser les limites des travaux existants et, ainsi, apporter un éclairage nouveau sur cette question.

Cette introduction est composée de trois parties. Dans un premier temps, nous dressons une vue d'ensemble de la théorie des zones monétaires optimales afin de déterminer les coûts et bénéfices de l'intégration monétaire. Dans un second temps, nous présentons un état de l'art des travaux empiriques débattant de cette question, ainsi qu'un bref aperçu de l'approche endogène des zones monétaires optimales. La troisième partie analyse les différentes options s'offrant à ces pays en matière de régime de change. Enfin, la dernière partie propose un résumé des contributions apportées dans cette thèse.

0.1 Les coûts et avantages de l'intégration monétaire

La question de l'intégration monétaire en Asie est largement débattue depuis la crise de 1997-98. L'accent a été posé sur la nécessité de coopérer plus étroitement dans les domaines financier et monétaire afin d'améliorer le fonctionnement des marchés financiers et de se protéger contre d'éventuelles attaques spéculatives. Des efforts importants ont été entrepris pour améliorer la surveillance économique régionale (*the Economic Review and Policy Dialog* en 1999, ERPD), promouvoir le développement des marchés obligataires en monnaie locale (*Asian Bond Markets initiative* et *Asian Bond Funds* en 2003, ABMI et ABF) et établir un mécanisme d'échange de devises sous forme de *swaps* (*the Chiang Mai Initiative Multilateralized* en 2009, CMIM). Dans le même temps, les ministres des Finances ont convenu de la nécessité de coordonner les politiques de change en vue de parvenir à une plus grande stabilité des taux de change intra-régionaux et, ainsi, favoriser le commerce intra-régional, l'investissement et l'intégration monétaire. Lors de la 39$^{\text{ème}}$ réunion annuelle de la Banque Asiatique de Développement en 2006, une première étape a été initiée dans cette direction avec la proposition de créer une unité monétaire asiatique (*Asian Currency Unit*, ACU) pour surveiller le mouvement relatif des devises asiatiques. Plus récemment, la perspective d'une monnaie commune en Asie a été défendue par le Premier ministre Japonais lors du 15$^{\text{ème}}$ sommet de l'Association des Nations de l'Asie du Sud-Est (ASEAN, 23 octobre 2009).[1] Les arguments en faveur d'une intégration monétaire plus poussée en Asie sont généralement liés à la prépondérance des forces du marché, tels que le commerce et les investissements directs à l'étranger (IDE), encouragées par la constitution d'un réseau intense des échanges intra-branches et la formation de plusieurs accords de libre-échange (ALE).[2]

Bien que des progrès substantiels aient été accomplis dans tous ces domaines, les questions relatives à la coordination monétaire régionale et au choix du régime de change demeurent sans réponse. La région reste en effet caractérisée par une grande diversité des régimes de change. Par exemple, les deux pays dominants en Asie, à savoir le Japon et la Chine, ont respectivement un régime de flottement pur et un régime de change où le taux yuan/dollar reste étroitement contrôlé. Depuis la crise de 1997-98, plusieurs propositions visant à renforcer l'intégration des politiques de change ont vu le jour, comme l'adoption d'un panier de devises, le retour à un ancrage souple sur le dollar ou l'adoption d'une monnaie unique. Certaines études ont suggéré que le dollar serait mieux adapté pour garantir une coordination implicite des taux de change et la stabilité du niveau des prix à l'échelle régionale (McKinnon and Schnabl , 2004). La création d'un "bloc yen" a également été envisagée par Kwan (2001) après que les autorités japonaises aient annoncé leur volonté d'internationaliser le yen, alors que l'idée d'une zone monétaire centrée sur le yuan commence seulement à émerger (voir, par exemple, Park , 2010). Certains économistes ont également envisagé la possibilité d'un panier (commun ou individuel) composé de devises internationales (Williamson , 2005) ou régionales (Ogawa and Shimizu , 2006a).

Sur le plan empirique, le débat a essentiellement porté sur l'existence d'une éventuelle zone monétaire optimale (ZMO). La théorie des ZMO examine sous quelles conditions deux ou plusieurs pays peuvent avoir un intérêt à partager une même monnaie. Les arguments sont principalement basés sur la comparaison des coûts et avantages relatifs à la fixité et la flexibilité du taux de change.[3] Les avantages d'une monnaie commune se reflètent principalement au niveau microéconomique. Le principal avantage provient de l'élimination des coûts de transaction qui pèsent sur les décisions économiques et sapent le commerce intra-régional. L'adoption d'une monnaie commune implique également moins d'incertitude car les firmes exportatrices et importatrices ne sont plus exposées à la variabilité des changes.[4] Une monnaie commune peut ainsi produire des bénéfices indirects, en ce qu'elle facilite les échanges commerciaux et stimule l'intégration économique.[5] La fixité des taux de change peut également se justifier

1. L'ASEAN comprend l'Indonésie, la Malaisie, les Philippines, Singapour, la Thaïlande, Brunei, le Vietnam, le Laos, la Birmanie et le Cambodge. L'ASEAN-5 fait référence à l'Indonésie, la Malaisie, les Philippines, Singapour et la Thaïlande, alors que l'ASEAN+3 regroupe les pays de l'ASEAN, la Chine, le Japon et la Corée du Sud. Afin de s'affranchir des problèmes liés à la toponymie, les termes "pays asiatiques", "Asie de l'Est" ou "Asie" feront, dans cette thèse, le plus souvent référence aux pays de l'ASEAN-5 et, dans une moindre mesure, à la Chine, au Japon et à la Corée du Sud. En tant que membres originels et économies les plus avancées de l'ASEAN, ce sont les pays les plus susceptibles d'initier le processus d'intégration monétaire en Asie avec la Chine et le Japon.

2. L'instauration d'une zone de libre-échange entre les pays de l'ASEAN (AFTA), lancée en 1992 et entrée en vigueur en 2002, est le signe le plus visible d'une intégration plus étroite entre les pays de l'ASEAN. D'autres mécanismes impliquant la Chine, le Japon et la Corée du Sud ont été établis, tels que l'ALE ASEAN-Chine (ACFTA, 2010), l'ALE Corée du Sud-ASEAN (AKFTA, 2010) et l'accord de partenariat économique global ASEAN-Japon (AJCEP, 2007). La Chine, le Japon et la Corée du Sud examinent actuellement la possibilité d'un ALE trilatéral.

3. Comme il a été souligné par Volz (2010), les facteurs qui sous-tendent l'analyse du compromis entre change flottant/change fixe et change flottant/union monétaire sont analytiquement les mêmes. Cependant, il est intéressant de noter que seule l'adoption d'une monnaie unique implique tous les avantages de l'intégration monétaire puisque la possibilité d'une réévaluation demeure en change fixe.

4. Il convient de noter que l'absence de marchés financiers sophistiqués dans les pays émergents réduit les possibilités de se couvrir contre le risque de change.

5. Plusieurs travaux récents tels que ceux de Arize et al. (2000), Hayakawa and Kimura (2009), Baum and Caglayan (2010) et Chit et al. (2010) montrent que la volatilité des changes a un impact négatif et statistiquement significatif sur les flux commerciaux bilatéraux.

lorsque des pays étroitement liés disposent d'une forte aversion à l'égard de l'appréciation réelle de leur taux de change, en ce sens que leur développement économique dépend principalement de la promotion de leurs exportations et donc de leur compétitivité (voir, par exemple, Kenen and Meade , 2008). Les pays dont la croissance est principalement tirée par les exportations peuvent ainsi trouver un intérêt particulier à opter pour un change fixe. L'autre principal avantage d'une monnaie commune est la possibilité de voir diminuer le différentiel de prix entre les pays grâce à une transparence accrue. Une monnaie commune devrait ainsi accroître la convergence nominale au sein d'une zone monétaire.

Le coût d'adhésion à une union monétaire découle de l'incapacité à mener une politique monétaire indépendante. Sous un régime de change fixe et sans contrôle des capitaux, la politique monétaire est vouée au maintien d'une cible externe (le taux de change), ce qui est incompatible avec l'adoption d'une cible interne comme la stabilité des prix ou la mise en oeuvre de politiques contra-cycliques. [6] Par exemple, sous un régime de change flottant, la banque centrale peut initier une politique monétaire expansionniste afin de stabiliser la production face à un choc de demande négatif. Inversement, l'assouplissement de la politique monétaire est incompatible avec le choix d'un régime de change fixe car l'augmentation de la masse monétaire affaiblit la valeur externe de la monnaie. Dès lors, le maintien d'une parité fixe implique une intervention sur le marché des changes et la perte de réserve de change, susceptible d'éroder la crédibilité de l'ancrage.

En outre, l'abandon de la monnaie nationale implique une perte totale de souveraineté monétaire aux dépens d'une politique monétaire commune. Cela est potentiellement coûteux lorsque les pays subissent des chocs asymétriques car l'utilisation du taux de change, comme instrument d'ajustement, devient impossible. Il en découle que le coût d'abandon d'une politique monétaire autonome est inversement proportionnel au degré de symétrie des chocs macroéconomiques. Mundell (1961) considère l'exemple de deux pays, l'Est et l'Ouest, qui forment une union monétaire en fixant le taux de change entre leur monnaie. Il fait l'hypothèse d'un choc asymétrique qui se manifeste par un glissement de la demande de l'Est vers l'Ouest. Cela se traduit mécaniquement par une baisse de la production et de la demande de travail à l'Est, alors que le mécanisme inverse se produit à l'Ouest. Ainsi, l'ajustement aura lieu sous forme d'inflation à l'Ouest et de chômage à l'Est. Mundell pose la question suivante: par quels mécanismes l'équilibre macroéconomique peut être atteint sans que les deux pays ne supportent le coût d'ajustement lorsque le taux de change est fixe?

Mundell insiste sur deux mécanismes qui amènent automatiquement l'équilibre dans les deux pays. Le premier est la flexibilité des salaires. Lorsque les salaires sont flexibles, l'excès de demande de travail à l'Ouest induit une pression à la hausse sur les salaire et les prix. Cela conduit mécaniquement les consommateurs des deux pays à se détourner des biens produits à l'Ouest, au profit des biens produits à l'Est qui deviennent plus compétitifs. L'ajustement se poursuivra jusqu'à ce que la perturbation initiale ait été compensée. Le second mécanisme est la mobilité de la main-d'oeuvre. Les chômeurs à l'Est se déplacent vers l'Ouest, où il existe un excès de demande de travail, ce qui résout respectivement le problème du chômage et de l'inflation à l'Est et à l'Ouest. La théorie des ZMO, telle que conçue par Mundell, conduit à la conclusion que le taux de change peut être abandonné comme instrument d'ajustement lorsque les chocs sont symétriques ou lorsqu'il existe des mécanismes d'ajustement alternatifs. Sinon, le choix d'un régime de change flottant est préférable. Avec un taux de change flexible,

6. Néanmoins, l'ancrage nominal extérieur n'est pas nécessairement incompatible avec la stabilité des prix domestiques. Dans le cas d'une petite économie ouverte, la fluctuation du taux de change peut se transmettre très rapidement au taux d'inflation (voir Calvo and Reinhart , 2001). Un régime de change fixe peut permettre d'importer la crédibilité de la monnaie ancre et d'atteindre la stabilité des prix domestiques.

et si les salaires sont rigides, l'Est peut en effet baisser son taux d'intérêt afin de stimuler la demande globale par une dépréciation de sa monnaie et *vice versa* à l'Ouest.

La symétrie des chocs, la mobilité des facteurs et la flexibilité des salaires constituent quelques-uns des critères standards pour former une ZMO. Depuis l'article fondateur de Mundell (1961), la littérature s'est efforcée d'identifier quelles autres conditions pouvaient rendre le choix d'un change fixe plus désirable. Ces autres conditions sont l'ouverture commerciale et la diversification des structures productives.

Lorsque les économies atteignent un niveau élevé d'intégration commerciale, McKinnon (1963) remet en cause l'utilité du taux de change comme instrument d'ajustement face aux chocs. Si une dépréciation de la monnaie stimule la demande pour les produits domestiques, elle augmente également le prix des produits importés. Lorsque le degré d'ouverture est élevé, l'effet d'amortissement positif est ainsi entièrement compensé par l'augmentation des salaires et des coûts de production. McKinnon (1963) établit ainsi une relation négative entre le coût d'un change fixe et l'ouverture commerciale d'une économie car une politique monétaire indépendante joue un rôle plus limité lorsque les pays sont fortement ouverts au commerce international. Kenen (1969) souligne la nécessité pour un pays d'avoir un degré élevé de diversification de ses structures productives lorsque les autorités monétaires souhaitent adopter un régime de change fixe. Si les chocs touchent spécifiquement certaines industries, les économies dont les structures productives sont largement diversifiées peuvent renoncer à la flexibilité car la dépendance de l'économie à l'égard d'un seul secteur industriel est réduite. De la même façon, si les pays possèdent des structures similaires en termes de production et de commerce extérieur, ils sont susceptibles d'être affectés simultanément par des chocs extérieurs, réduisant ainsi l'occurrence des chocs asymétriques.

0.2 Les pays asiatiques remplissent-ils les conditions d'une ZMO? Un état de l'art.

Comme expliqué ci-dessus, lorsque les pays font face à des chocs symétriques, la nécessité d'établir des mécanismes alternatifs pour atteindre la stabilité macroéconomique est considérablement réduite. C'est pourquoi, la synchronisation des cycles économiques est un critère fondamental pour des pays souhaitant adopter une même monnaie. Concernant les pays asiatiques, la littérature n'ait que très peu concluante sur cette question car les résultats empiriques sont particulièrement sensibles au choix de l'approche économétrique.

Une approche fréquemment utilisée pour évaluer la corrélation des chocs au sein d'un groupe de pays est la méthodologie du VAR structurel (SVAR) proposée par Blanchard and Quah (1989). Cette méthode est basée sur l'imposition de restrictions de long terme, de manière à pouvoir identifier la dynamique des chocs affectant la production réelle de façon permanente et transitoire.[7] Bayoumi and Eichengreen (1994) constatent que les chocs d'offre sont symétriques d'une part, entre Singapour, la Malaisie, l'Indonésie et Hong Kong, et d'autre part, entre le Japon, Taïwan et la Corée du Sud. Chow and Kim (2003) montrent que les pays d'Asie du Sud-Est restent fortement influencés par des chocs spécifiques contrairement aux pays de l'Union Européenne dont la dynamique du cycle économique est principalement affectée par des chocs régionaux. Ahn et al. (2006) trouvent un groupe de sept pays

7. D'un point de vue théorique, les chocs de demande n'ont qu'un impact transitoire sur le produit à long terme tandis que les chocs d'offre influencent la production d'une manière permanente.

(Indonésie, Malaisie, Singapour, Thaïlande, Hong Kong, Corée du Sud et Taïwan) qualifié pour une ZMO. L'analyse menée par Huang and Guo (2006) suggère qu'il peut être bénéfique pour Hong Kong, l'Indonésie, la Malaisie, Singapour et la Thaïlande de former une union monétaire car les perturbations affectant ces économies sont corrélées positivement. Inversement, Kim (2007) constate que les chocs macroéconomiques restent hétérogènes en Asie du Sud-Est, ce qui amène l'auteur à considérer qu'une union monétaire serait prématurée pour ces pays. Tenant compte de la dynamique des chocs à travers le temps, Zhang and Sato (2008) suggèrent que les économies de la Grande Chine (Chine continentale, Hong Kong et Taïwan) se posent comme des candidats sérieux à l'adoption d'une monnaie unique. Obiyathulla (2008) examine la faisabilité d'une zone monétaire entre quatorze pays d'Asie et constate que, prise dans son ensemble, la région ne forme pas une ZMO, bien que plusieurs groupes séparés se posent en candidats potentiels. Enfin, Genberg and Siklos (2010) échouent à identifier un groupe de pays pour lesquels les chocs sont sans ambiguïté fortement corrélés, alors que Lee and Koh (2012) sont favorables à une intégration monétaire plus poussée entre les pays de l'ASEAN. En outre, les auteurs soulignent que le degré de symétrie des chocs a augmenté après la crise financière de 1997-98.

Cependant, l'approche par le modèle SVAR reste ouverte à de nombreuses critiques. Tout d'abord, ce type de modèle implique le plus souvent l'utilisation de variables en différence première, éliminant ainsi toute information relative à l'existence d'une relation d'équilibre de long terme. Comme l'ont souligné Lee and Azali (2012), on peut également s'attendre à ce que les chocs de demande soient persistants, sinon permanents, en particulier lorsque les économies ont des taux de chômage élevés. Dans le cas où un choc de demande affecte la technologie et la productivité, un déplacement de la courbe de demande agrégée induit un changement du niveau potentiel de la production, ce qui, à son tour, déplace la courbe d'offre agrégée de long terme. [8] Les hypothèses d'orthogonalité des chocs d'offre et de demande, ainsi que de l'effet transitoire des chocs de demande ne sont donc pas incontestables (Cover et al. , 2006).

Une approche complémentaire est le concept de la parité des pouvoirs d'achats généralisée (G-PPP) introduit par Enders and Hurns (1994) afin de surmonter l'échec empirique de la parité des pouvoirs d'achats (PPP). Les auteurs soutiennent que les fondamentaux macroéconomiques qui influent sur l'évolution des taux de change réels ne sont pas stationnaires (ou $I(1)$) et, par conséquent, impliquent des mouvements permanents de ces derniers. Partant de cette observation, la possibilité qu'ils partagent une tendance commune de long terme -lorsque les variables qui sous-tendent leur évolution sont similaires- devient tout à fait envisageable. Dans ce cas, les économies sont soumises à des chocs communs et il devient optimal pour elles de former une union monétaire. Wilson and Choy (2007) testent la théorie de la G-PPP pour déterminer si les taux de change réels de l'ASEAN-5 -à la fois en termes de yen et de dollar- partagent une tendance stochastique commune de long terme. Leurs résultats ne concluent pas en faveur d'une ZMO avec le Japon et les États-Unis même après la crise asiatique. Mishra and Sharma (2010) trouvent des résultats opposés pour sept pays de l'ASEAN et l'Inde. Ils suggèrent que ces pays forment une ZMO avec le Japon ou les États-Unis qui sont leurs principaux partenaires commerciaux. Selon les auteurs, le yen et dollar ont une importance comparable pour les pays d'Asie, et un ancrage commun sur un panier composé des deux monnaies serait plus optimal qu'un ancrage unilatéral. Ces résultats confirment les conclusions obtenues par Choudhry (2005) qui ne trouve aucune différence dans

8. Cela suppose que des changements dans la fonction d'offre d'une économie ne soient pas indépendants des changements liés à la demande. Par exemple, un changement dans les comportements des consommateurs peut induire un changement dans l'utilisation des *inputs*. Il en résulterait une réorganisation des compétences au sein de l'industrie et à une allocation des ressources davantage tournée vers la recherche et le développement (voir Stadler , 1990).

le choix de la monnaie d'ancrage. De la même manière, Ogawa and Kawasaki (2008) affirment que les pays de l'ASEAN+3 forment une ZMO, ouvrant ainsi la voie à l'adoption d'une politique de change commune autour d'un panier de devises comprenant le dollar, l'euro et le yen. Sun and Simons (2011) étudient l'éventualité d'une relation de long terme entre les taux de change effectifs réels. Ils trouvent un "pentagone" composé de la Corée du Sud, des Philippines, de la Thaïlande, de l'Indonésie et de la Malaisie, susceptibles de former une union monétaire. Enfin, Rangkakulnuwat et al. (2010) analysent la G-PPP où l'offre de monnaie, la production et le taux d'intérêt japonais sont les variables supposées guider l'évolution des taux de change réels à long terme. Les auteurs suggèrent que la Corée du Sud, la Malaisie, les Philippines, Singapour et la Thaïlande peuvent constituer une zone monétaire où le yen serait la monnaie d'ancrage.

La littérature sur la synchronisation des chocs en Asie compte beaucoup d'autres travaux empiriques dont les approches se veulent également originales. Citons les travaux de Girardin (2004) et Girardin (2005) dont l'approche consiste à supposer que cette synchronisation dépend de la phase du cycle de croissance. Utilisant une approche Markov-switching VAR (MS-VAR), ses résultats ne soutiennent pas la possibilité d'un système de change centré autour du yen car les différentes phases du cycle économique ne sont pas toutes synchrones entre le Japon et les pays asiatiques. Par ailleurs, selon l'auteur, la Chine serait un meilleur candidat dans le rôle du pays ancre. Par le biais de la cointégration, Sato and Zhang (2006) constatent que certains groupes de pays partagent -à la fois sur le court et le long terme- des mouvements synchrones de leur production réelle. Les auteurs revendiquent la création d'une union monétaire entre Singapour, la Thaïlande et l'Indonésie d'une part, et Hong Kong, la Corée du Sud, la Chine, le Japon et Taïwan d'autre part. Sato et al. (2009) montrent que les pays de l'ASEAN ne constituent pas un groupe optimal pour former une union monétaire à moins que le Japon ne soit inclus. Socorro Gochoco-Bautista (2008) estime la composante commune relative à la croissance de la production industrielle des pays d'Asie, et analyse si ce *benchmark* est statistiquement significatif dans l'explication des fluctuations de la production de chaque pays. Les conclusions de l'auteur indiquent qu'il existe un facteur régional important qui explique le mouvement des cycles économiques. Moneta and Rüffer (2009) utilisent un modèle espace-état où cette composante commune est modélisée selon un facteur inobservable commun à toutes les séries. L'auteur trouve une dynamique commune et significative dans la croissance des pays asiatiques. Allegret and Essaadi (2011) effectuent une analyse spectrale des cycles économiques en Asie et détectent la présence d'un cycle commun après la crise de 1997-98. Les auteurs suggèrent que les pays d'Asie constituent une ZMO. À partir d'une approche bayésienne des modèles espace-état, Lee and Azali (2012) examinent dans quelles mesures le produit des pays asiatiques est influencé par des chocs globaux, régionaux ou locaux. Les auteurs constatent un rôle croissant du facteur régional, bien que les facteurs spécifiques restent importants dans l'évolution des cycles. Enfin, Nguyen (2010) utilise une approche similaire et constate que les pays asiatiques sont des candidats moins plausibles pour une union monétaire que les pays européens.

Si l'adoption d'une monnaie commune est généralement soumise à l'examen préalable des critères discutés ci-dessus, la théorie des ZMO ignore la dimension endogène liée au processus d'intégration monétaire. En effet, les pays dont les chocs macroéconomiques ne sont pas suffisamment corrélés peuvent tout de même trouver un intérêt à former une union monétaire car l'adoption d'une monnaie commune favorisera leur intégration commerciale et la synchronisation des cycles économiques. En d'autres termes, l'approche endogène des ZMO suggère qu'un pays est davantage susceptible de satisfaire les critères d'optimalité *ex-post* plutôt que *ex-ante* (Frankel and Rose , 1998).

Théoriquement, l'effet de l'intégration commerciale sur la synchronisation des cycles économiques est ambigu. Lorsque la plupart des échanges sont intra-branches, l'intensification des flux commerciaux devrait diminuer l'occurrence des chocs asymétriques. Dans une telle configuration, les pays achètent et vendent les mêmes catégories de produits, ce qui devrait se traduire par une augmentation de la co-variance des chocs de demande. Inversement, l'intensification du commerce peut aussi conduire à une plus grande concentration des activités industrielles dans une région, diminuant ainsi la synchronisation des cycles économiques (Krugman , 1991). Cette question a été empiriquement étudiée par Frankel and Rose (1998). Les auteurs examinent la relation entre l'intégration commerciale et la synchronisation des cycles économiques pour un panel de 210 corrélations, et constatent que les pays dont les liens commerciaux sont importants possèdent une corrélation plus forte entre leur cycle économique. Les auteurs en déduisent que la formation d'une union monétaire favorise l'intégration commerciale qui conduit elle-même à synchroniser davantage les cycles économiques. Ainsi, à mesure que l'intégration commerciale se renforce en Asie, une synchronisation plus étroite entre les cycles économiques peut être attendue. Par exemple, Shin and Wang (2003) observent que le commerce intra-branche, ou le volume du commerce lui-même, conduit à une meilleure synchronisation de douze cycles asiatiques. Des résultats similaires sont obtenus par Shin and Sohn (2006) et Rana (2007). Dans cette perspective, l'union monétaire devrait être considérée non pas comme une finalité mais plutôt comme un moyen de renforcer l'intégration économique. Rose (2000) montre en effet que deux pays partageant la même monnaie ont tendance à commercer trois fois plus. Rose and van Wincoop (2001) montrent que l'adoption d'une monnaie unique en Europe tendrait à augmenter le commerce européen de plus de 50%. Glick and Rose (2002) utilisent des données de panel pour un ensemble couvrant 217 pays et trouvent que l'appartenance à une même zone monétaire double la taille des flux commerciaux. Enfin, Shirono (2008) utilise un modèle de gravité micro-fondé et estime que l'adoption d'une monnaie commune en Asie pourrait doubler les échanges commerciaux bilatéraux dans la région.

0.3 Quel régime de change pour le futur?

Si l'approche par les ZMO, qu'elle soit statique ou dynamique, offre un cadre théorique général pour se prononcer sur la faisabilité d'une union monétaire, elle ne tient pas compte de certaines spécificités propres aux pays asiatiques. D'une part, l'intégration monétaire est un processus qui nécessite une forte adhésion politique et un degré élevé de coopération. Aujourd'hui, un tel engagement fait encore défaut en Asie (voir, par exemple, Plummer , 2006, Volz , 2006, Chey , 2009). D'autre part, les arguments qui justifient la nécessité pour les pays asiatiques de préserver une stabilité relative contre le dollar ne peuvent être appréhendés par la théorie des ZMO. Par ailleurs, il n'existe pas encore de monnaies régionales qui pourraient agir comme un point d'ancrage dans la coopération et la coordination des politiques de change. Même si le yen peut théoriquement remplir cette fonction, son rôle demeure limité en Asie. Enfin, le manque de maturité des marchés financiers chinois ainsi que la non-convertibilité du yuan exclut actuellement la possibilité d'un "bloc yuan". Comme nous allons le voir ci-dessous, cela laisse au moins trois alternatives pour renforcer la stabilité des taux de change intra-régionaux: la création d'un "bloc dollar", un ancrage sur un panier composé des principales devises internationales, ou la création d'une monnaie asiatique comme l'ACU.

Compte tenu de leur forte ouverture sur le reste du monde et du niveau élevé d'intégration commerciale, la volatilité exacerbée des taux de change intra-régionaux peut avoir un effet très déstabilisant sur les économies asiatiques. L'adoption d'un flottement pur dans toute la région ne semble donc pas compatible avec le degré élevé d'interdépendance économique atteint par ces pays. Par ailleurs, même si les pays asiatiques semblent avoir introduit après la crise Asiatique davantage de flexibilité dans la gestion de leur taux de change, il n'en demeure par moins que ceux-ci restent étroitement contrôlés par les autorités monétaires. Parmi d'autres, McKinnon and Schnabl (2004) montrent que tous les pays ont abandonné leur ancrage durant la crise asiatique, à l'exception de la Chine et de Hong Kong. Cependant, après la crise, le poids du dollar semble atteindre selon les auteurs des niveaux comparables à ceux observés avant la crise dans la plupart des pays (sauf en Indonésie et aux Philippines). Ainsi, les pays asiatiques interviendraient fréquemment sur le marché des changes malgré l'adoption officielle d'un régime de change flexible, ce qui traduirait une forte aversion a l'égard d'une volatilité exacerbée de leur devise. [9]

Pour McKinnon and Schnabl (2004), la résurrection de "l'étalon dollar" en Asie est motivé par le fait que cette devise continue de jouer un rôle important dans les transactions commerciales au sein de la région, tandis que les États-Unis restent un partenaire commercial de premier plan en Asie. Par ailleurs, de nombreux pays asiatiques ne peuvent emprunter sur les marchés internationaux avec leur propre devise, amenant les institutions financières ainsi que les agents privés à cumuler un déséquilibre de devises (*currency mismatch*). Ce déséquilibre augmente la vulnérabilité des systèmes financiers des pays débiteurs face à une éventuelle dépréciation du taux de change domestique. En revanche, depuis la fin des années 1990, de nombreux pays Asiatiques tendent à dégager des excédents de leur compte courant avec les États-Unis. Il en résulte que l'accumulation d'un stock de créances libellées en dollar, par les autorités monétaires et les agents privés, fait peser le risque d'un *rush* auto-entretenu sur la monnaie nationale qui conduirait à l'appréciation de leur devise et à une baisse de la valeur de ces actifs. [10] Le maintien d'une parité relativement fixe sur le dollar constitue donc un mécanisme d'assurance implicite contre une trop forte volatilité dans la valeur des portefeuilles domestiques. Enfin, un ancrage commun sur le dollar peut fournir aux pays asiatiques un mécanisme de coordination implicite des politiques de change garantissant, dans une certaine mesure, la stabilité des taux de change intra-régionaux et des prix à l'échelle régionale.

Le désir de maintenir sous-évaluée leur devise peut donc également justifier ce choix lorsque l'on considère le besoin pour ces pays de préserver leur compétitivité, dans la poursuite d'une stratégie de croissance tirée par les exportations. Coudert et al. (2013) montrent en effet que les pays asiatiques ont tendance à déserrer leur ancrage sur le dollar lorsque celui s'apprécie au-dessus d'un certain seuil, et cela dans le but de préserver leur compétitivité face à leurs partenaires commerciaux. Pour Ito et al. (1998) et Ogawa and Ito (2002), cet aspect doit être assimilé à un défaut de coordination dans le choix d'un système de change optimal au sein de la région car aucun pays n'aurait intérêt à abandonner son ancrage sur le dollar, aussi longtemps que les autres pays continuent de stabiliser leur taux de change contre celui-ci.

Un ancrage commun sur le dollar ne saurait toutefois éliminer la vulnérabilité des monnaies asiatiques face aux fluctuations des principales devises internationales, telles que le yen ou l'euro. Il est généralement admis que la forte dépréciation du yen au cours de la seconde moitié des années 1990 sem-

9. Calvo and Reinhart (2002) qualifient cette situation de "peur du flottement" (*fear of floating*). Les raisons pour lesquelles les pays émergents ne sont pas disposés à tolérer une trop forte variation de leur taux de change sont aussi analysés par Hausmann et al. (2001). Les arguments avancés sont ceux relatifs au degré du *pass-through*, la capacité pour un pays d'emprunter des capitaux sur les marchés internationaux dans sa propre devise et, enfin, la volatilité élevée des taux d'intérêt qui limite les gains d'indépendance monétaire liés à la flexibilité.

10. Cette crainte est également exacerbée par le fait que les devises des pays asiatiques demeurent sous-évaluées, appelant tôt ou tard à un ajustement de ces dernières afin de résorber le déficit courant des États-Unis. McKinnon and Schnabl (2004) parlent de *conflicted virtue* pour caractériser cette situation.

ble avoir provoqué un choc externe qui amplifia considérablement la crise de 1997-98 (Kwan , 2001; McKinnon and Schnabl , 2004; Moneta and Rüffer , 2009). C'est pourquoi la création d'un système de change asiatique centré sur le yen a également été envisagé (Kwan , 2001). Cette option est basée sur l'idée que les pays asiatiques doivent stabiliser leur taux de change contre le yen et utiliser celui-ci comme une monnaie de référence dans les transactions commerciales et financières. Cette stabilisation contre le yen peut être souhaitable compte tenu de l'étroite interdépendance économique entre le Japon et ses partenaires asiatiques en termes de commerce et d'IDE. Lorsque le yen se déprécie par rapport au dollar, les pays asiatiques qui sont en compétition avec le Japon sur les marchés internationaux et régionaux, voient leur compétitivité se dégrader et leurs exportations diminuer. De même, les IDE à destination de ces pays ont tendance à ralentir car les coûts de production et d'investissement deviennent relativement plus onéreux qu'au Japon. En d'autres termes, une dépréciation du yen est principalement associée à un ralentissement de l'économie de ces pays, alors qu'une appréciation du yen stimule leur croissance. Cependant, un régime de change centré sur le yen reste difficilement envisageable. Notamment, le yen ne peut jouer un rôle central dans la coopération monétaire régionale malgré la volonté affichée des autorités japonaises de promouvoir l'internationalisation du yen. Suite à l'éclatement de la bulle spéculative survenue au Japon à la fin des années 1980, ce dernier a connu une lente dégradation de sa situation économique, tandis que la croissance des autres pays d'Asie a été fulgurante jusqu'à la crise de 1997-98. Par conséquent, l'utilisation du yen dans la politique de change de ces pays est restée limitée. De plus, même si le Japon reste un important partenaire commercial, son leadership est désormais concurrencé par celui de la Chine. Shirono (2009) observe notamment qu'une union monétaire avec la Chine tend à générer plus de bien-être qu'une union monétaire avec le Japon.

Si la stabilisation des taux de change intra-régionaux est souhaitable, aussi longtemps que les principales devises internationales fluctueront entre-elles, un ancrage exclusif sur le yen, le dollar ou l'euro ne peut être envisagé comme une solution crédible et souhaitable pour renforcer la coordination des politiques de change en Asie. C'est pourquoi, après la crise de 1997-98, beaucoup d'observateur ont suggéré l'adoption d'un ancrage sur un panier composé de devises externes et/ou régionales. Par exemple, la stabilisation des monnaies asiatiques vis-à-vis du yen pourrait atteinte grâce à un ancrage sur un panier de devises à l'intérieur duquel il aurait un poids prépondérant. Le poids optimal qui viserait à stabiliser la production ainsi que la balance des paiements serait plus important pour les pays qui sont en concurrence avec le Japon sur les marchés internationaux (les nouveaux pays industrialisés, NPIs), et plus faible pour les autres pays (l'ASEAN et la Chine). Selon Kwan (2001), ce panier de devises serait une première étape vers une union monétaire entre le Japon et les NPIs. Les pays les moins avancés pourraient adhérer à l'union monétaire après une période de transition au cours de laquelle ils augmenteraient progressivement le poids du yen. Plusieurs économistes suggèrent d'augmenter sensiblement le poids du yen dans la gestion de la politique de change, sans revendiquer toutefois la création d'un "bloc yen". Prenant en compte l'impact économique induit par la forte volatilité du taux de change dollar/yen, Kawai (2002) propose que les pays asiatiques stabilisent leur taux de change vis-à-vis d'un panier de devises où le dollar, le yen et l'euro joueraient un rôle égal. Bénassy-Quéré (1999) constate que le poids optimal du yen est inférieur à celui observé dans la réalité, en dépit de l'importance de ce dernier dans la composition de la dette extérieure. Cela amène l'auteur à se prononcer en faveur d'une politique de change centrée sur un panier de devises où le poids du yen serait plus élevé. L'adoption d'un panier de devises dont la composition serait équitablement répartie entre le yen, le dollar et éventuellement, l'euro, a également été suggérée par Ito et al. (1998), Williamson (1999, 2001), Rajan (2002), Bird and Rajan (2002).

En général, l'objectif pour un pays d'ancrer son taux de change sur un panier de devises est de stabiliser son taux de change effectif nominal (NEER) afin de réduire la volatilité de son taux de change effectif réel (REER) et, ainsi, préserver sa compétitivité vis-à-vis de ses principaux partenaires commerciaux. Le REER est souvent considéré comme un indicateur de compétitivité car son augmentation indique, en moyenne, une appréciation réelle de la monnaie domestique par rapport aux devises des autres pays qui entrent dans la composition du taux de change effectif. Ce choix peut se justifier lorsque la répartition géographique des échanges extérieurs n'est pas polarisé sur un ensemble restreint de partenaires (Wilson et al. , 2007). Cela est notamment le cas des pays asiatiques dont la composition du commerce extérieur reste très diversifiée avec la Chine, le Japon, les États-Unis et l'Union Européenne comme principaux partenaires. Ainsi, la stabilisation du NEER pourrait être plus optimale que le maintien d'une parité vis-à-vis d'une seule monnaie. Cette stratégie aurait l'avantage d'accroître la stabilité des changes intra-régionaux car la distribution du commerce extérieur est relativement identique au sein de la région. Les paniers de chaque pays incluraient en effet les mêmes devises avec des poids relativement similaires. Par ailleurs, l'adoption simultanée d'un régime qui viserait à piloter chaque taux de change contre un panier commun de devises pourrait constituer un mécanisme de coopération des politiques de change à l'échelle régionale. Enfin, un panier commun offrirait aux pays asiatiques l'avantage de laisser flotter conjointement leur monnaie face au dollar, au yen ou à l'euro.

Wilson et al. (2007) soulignent que l'adoption d'un panier individuel, dont les pondérations et le choix des devises seraient adaptées à chaque pays, implique un compromis entre le gain attribuable à la réduction de la volatilité du NEER et le coût associé à l'augmentation des fluctuations entre les taux de change intra-régionaux. Les auteurs simulent une analyse contrefactuelle basée sur une comparaison entre le régime de change historique et différents paniers de devises. Ils constatent que l'ancrage sur un panier individuel minimise significativement la volatilité du NEER pour tous les pays asiatiques. Ce gain est au moins aussi important que le celui induit par l'adoption d'un panier commun. Ohno (1999) constate également une faible différence en termes de stabilité du taux de change. Cependant, Williamson (1998) observe que l'adoption d'un panier individuel implique une forte instabilité des taux de change par rapport au dollar, augmentant ainsi les fluctuations des taux de change intra-régionaux. Inversement, la stabilisation du taux de change vis-à-vis d'un panier commun offrirait la possibilité de supprimer définitivement la volatilité des taux de change intra-régionaux. Plus récemment, Williamson (2009) obtient des conclusions similaires pour neuf pays asiatiques. L'auteur montre que le gain induit par l'adoption d'un panier commun serait supérieur au gain induit par l'adoption d'un panier individuel. Williamson (1998) préconise également l'adoption d'une bande de fluctuation mobile autour de la valeur du panier ciblé afin de limiter les interventions sur le marché des changes et la possibilité d'une attaque spéculative.

La dernière option pouvant être envisagée consisterait en la création d'une unité monétaire asiatique nommée ACU. En Décembre 2006, la Banque de Développement Asiatique a effectué un premier pas dans cette direction en annonçant vouloir créer un étalon de référence pour surveiller les mouvements dans la valeur relative des devises asiatiques. Depuis lors, la création d'un mécanisme de coordination dans lequel l'ACU jouerait un rôle central est considéré comme une option réaliste pour les années à venir. L'ACU se définit comme un panier des devises de l'ASEAN+3, pondéré selon le poids respectif de chaque pays dans le PIB et le commerce régional. Dans cet esprit, Ogawa and Shimizu (2006a, 2006b) ont calculé l'écart de chaque devise asiatique par rapport à ce *benchmark* et ont constaté un désalignement significatif des monnaies asiatiques. Ils interprètent leur résultat comme une illustration du manque de coordination des politiques de change au sein de la région. En outre, les auteurs comparent leur résultat avec le panier commun G-3 défendu par Williamson (2005), et montrent qu'un système d'ancrage sur l'ACU amène à une plus grande stabilité du NEER pour l'Indonésie, les Philippines, la Corée du Sud et la Thaïlande.

Toutefois, la création de l'ACU n'exclut pas nécessairement la possibilité d'un ancrage sur un panier de devises. Une stratégie possible pour les pays asiatiques consisterait dans un premier temps à opérer un mouvement collectif de leur régime de change actuel, vers un régime de change centré sur un panier individuel de devises. Chaque pays adapterait ainsi les pondérations au sein de son propre panier afin de stabiliser son NEER. Dans le même temps, l'introduction de l'ACU -comme indicateur statistique- permettrait de contrôler le mouvement collectif des monnaies asiatiques par rapport au dollar, à l'euro et au yen si ce dernier n'est pas inclus dans l'ACU. Selon Kawai (2009), l'ACU peut jouer un rôle crucial dans le développement de nouveaux instruments financiers négociables et, plus généralement, dans la promotion des marchés de capitaux régionaux. Comme Ogawa and Shimizu (2009) l'ont souligné, l'émission d'obligations libellées en ACU peut réduire les coûts d'emprunt de capitaux étrangers ainsi que le risque de change pour les investisseurs internationaux. Cela permettrait aux autorités régionales d'apporter une solution efficace à la nécessité de se doter d'un marché obligataire régional plus liquide et profond, en vue de trouver des sources alternatives de financement. De même, Eichengreen (2006) considère que l'ACU pourrait agir comme une monnaie parallèle qui circulerait en même temps que les autres monnaies nationales. Au fur et à mesure que les taux de change se stabiliseraient vis-à-vis du panier régional, et donc les uns par rapport aux autres, l'ACU pourrait être utilisé comme un véhicule encourageant le commerce, l'investissement et les transactions financières au sein de la région. Enfin, à un stade plus avancé, l'ACU pourrait jouer un rôle officiel semblable à celui joué par *l'European Currency Unit* (ECU) au sein du Système Monétaire Européen (SME). Par exemple, l'ACU pourrait être mis en circulation comme unité de compte, moyen d'échange ou réserve de valeur, tandis que les autorités pourraient décider de stabiliser officiellement leur taux de change contre l'ACU (Kawai , 2009).

0.4 Contributions de cette thèse

Dans cette thèse, nous proposons quatre contributions originales à l'étude de l'intégration monétaire et financière des pays d'Asie du Sud-Est. Le troisième chapitre a été publié dans la revue *Journal of International Financial Markets, Institutions and Money*. Les trois autres chapitres sont actuellement soumis dans des revues de rang international. Par le biais de techniques économétriques récentes et appropriées, nous nous efforçons tout au long de cette étude d'apporter un éclairage nouveau sur quatre problématiques fondamentales: la coordination des politiques de change, la synchronisation des cycles des affaires, la cointégration des taux de change réels et le degré d'intégration financière. Cette thèse se compose essentiellement de travaux empiriques dont la finalité est la recommandation de politiques économiques visant à renforcer le processus d'intégration monétaire et financière de ces pays.

Contribution du chapitre 1 - Ce premier chapitre examine la coordination des politiques de change en Asie (ASEAN-5 et Corée du sud) autour d'une unité monétaire synthétique nommée ACU et définie comme un panier des principales devises asiatiques. Nous calculons cette unité monétaire synthétique pondérée selon le poids respectif de chaque devise/pays dans le commerce et le PIB régional. Nous proposons d'estimer un modèle VAR afin d'étendre le modèle empirique de Frankel and Wei (1994), et cela, dans le but d'examiner la sensibilité relative des devises asiatiques face aux chocs structurels simulés sur le dollar, l'euro et l'ACU. Nous montrons que la stabilité des monnaies asiatiques autour de l'ACU reste faible avant 2006. Nous l'expliquons par la prépondérance du dollar au sein de la région après la crise. En revanche, l'influence du dollar décroît après 2006, alors que le poids de l'ACU, lui, gagne en importance dans la gestion des politiques de change asiatiques. Ce résultat met en évidence la volonté des pays asiatiques de se détourner d'une politique de change exclusivement centrée sur le dollar vers une politique plus flexible, avec un panier de devises comme point d'ancrage.

Contribution du chapitre 2 - Le deuxième chapitre met l'accent sur la synchronisation entre les cycles des affaires des pays de l'ASEAN-5. Nous cherchons à établir si les corrélations entre les cycles des affaires dépendent du régime du cycle (phases de contraction et de récession). Nous discriminons également les corrélations entre source régionale et globale afin d'établir si ces corrélations proviennent de la transmission des cycles à l'échelle régionale (entre les pays de l'ASEAN-5) ou si elles sont le fruit de la dynamique des cycles globaux (Chine, Japon, USA) supposés jouer un rôle important dans la synchronisation au sein de la région. Nous utilisons un modèle Markov-switching dont les probabilités de transition varient dans le temps (Filardo , 1994) afin de modéliser la non-linéarité et la dynamique des corrélations. Nous montrons que la corrélation entre les cycles est plus forte durant les phases de contraction mais que la dynamique d'ajustement est propre à chaque pays. Par ailleurs, contrairement aux cycles globaux, certains cycles des affaires asiatiques contiennent des informations pertinentes pour prédire les changements de régime des autres pays, ce qui tendrait à prouver que l'ASEAN-5 est découplé du reste du monde. Dans l'ensemble, nos résultats montrent qu'une union monétaire entre les pays de l'ASEAN-5 reste prématurée.

Contribution du chapitre 3 - Ce troisième chapitre examine la théorie de la parité de pouvoir d'achat généralisé (G-PPP) (Enders and Hurns , 1994, 1997) entre les taux de change réels de l'ASEAN-5. Plus spécifiquement, nous cherchons à identifier si ces taux de change réels partagent un équilibre de long terme et dans quelle mesure les écarts à l'équilibre sont persistants. Nous utilisons les estimateurs de Nielsen and Shimotsu (2007) et Shimotsu (2012) afin de procéder à l'analyse du rang et de la régression du système de cointégration. Nous identifions plusieurs relations de cointégration fractionnaire dites “faibles”. En conséquence, les résidus du système de cointégration sont très persistants mais *mean-reverting*. Nos résultats contrastent avec ceux des études antérieures qui limitent leur enquête à la cointégration traditionnelle $I(1)/I(0)$. Contrairement aux résultats existants, nous montrons en effet que les taux de change réels sont liés par un processus à mémoire longue. Ainsi, nos résultats supportent l'idée d'une plus forte intégration monétaire entre différents sous-groupes de pays car leurs taux de change réels partagent un équilibre de long terme. En revanche, une union monétaire comprenant tous les pays de l'ASEAN-5 semble actuellement compromise. Dans la perspective d'une future union monétaire, nous montrons également qu'un ancrage régional sur le dollar serait moins désirable qu'un ancrage sur un panier de devises composé du yen et du dollar.

Contribution du chapitre 4 - Dans ce dernier chapitre, nous examinons le degré d'intégration entre les marchés boursiers asiatiques (l'ASEAN-5, le Japon et Hong Kong) en nous focalisant sur le comouvement des volatilités observées. Nous étudions également la relation entre ces marchés asiatiques et les États-Unis. Dans cette optique, nous utilisons un modèle de cointégration fractionnaire afin de capturer la part de la volatilité dont l'origine est commune. Notre méthodologie est capable de couvrir à la fois les régions stationnaires et non-stationnaires de l'espace des paramètres. Nos résultats montrent que la volatilité sur les marchés boursiers de Singapour, de Hong Kong, des États-Unis, et dans une moindre mesure, du Japon, partagent une tendance stochastique commune. Ce résultat est cohérent avec l'idée selon laquelle le processus de long terme qui sous-tend la dynamique de la volatilité des marchés boursiers internationaux est guidé par des facteurs globaux d'intégration. En revanche, les marchés boursiers des pays émergents ne sont ni cointégrés entre eux, ni avec les marchés internationaux. De ce point de vue, ces marchés financiers apparaissent peu intégrés tant au niveau régional qu'au niveau global, limitant ainsi les possibilités pour ces pays de lisser leur consommation et partager le risque macroéconomique face aux chocs spécifiques.

Chapter

1 Assessing Asian Exchange Rates Co-ordination under Regional Currency Basket System

1.1 Introduction

This empirical paper gives new evidence concerning the coordination of exchange rate policies in Asia, by examining the degree of intra-regional exchange rate stability around the Asian Currency Unit (ACU), the US dollar and the euro. The 1997-98 currency crisis highlighted the close economic inter-dependence among the Asian countries. This leads the regional authorities to agree upon the need to promote a collective arrangement in order to stabilize their exchange rates and foster monetary policy coordination.[1] The market-driven integration, trough trade and foreign direct investment, is actually oriented toward the adoption of a common currency basket system. Prior to the crisis, the common US dollar pegging allowed implicit exchange rate stabilization (McKinnon , 1998; McKinnon and Schnabl , 2004). However, the crisis emphasized the fragility of rigid exchange rate arrangements notably for countries with a diversified trade pattern. The yen's depreciation against the US dollar from mid-1990 is particularly illustrative. The Asian currencies that were linked to the US dollar became overvalued and vulnerable to the volatility of the yen/dollar exchange rate. This third-currency effect is believed to be some of the main causes of the 1997-98 crisis because Asian export competitiveness declined against Japanese products in regional and third markets as the yen depreciated (Kwan , 2001; Bird and Rajan , 2002).[2] Since then, it is commonly assumed that an exclusive anchor to the US dollar (or the yen and the euro) is neither a credible nor a desirable solution for the future.

Recognizing this, most of the crisis-hit countries have officially abandoned the US dollar as an unilateral anchor since the crisis. The exchange rate policies within the region have evolved considerably and the coordination of exchange rates appears to be difficult to achieve at the regional level. After the crisis, some countries have adopted a single currency peg (Hong Kong but also China and Malaysia up to July 2005), whereas other countries have officially operated flexible exchange rate regimes (currency baskets, crawling bands ect.). Although a full-fledged monetary union is regarded as unrealistic, at least in the short term, numerous recent studies advocate for the adoption of a gradual step approach starting with informal forms of policy coordination. Williamson (2005) proposes a common basket peg (BBC) composed of the US dollar, the yen and the euro, for nine countries (China, Thailand, Philippines, Singapore, Taiwan, South Korea, Malaysia, Indonesia and Hong Kong). Kawai (2002), Mori et al.

1. The prospect of launching a single currency was put forward by the Japanese prime minister on 23 October 2009 during the 15th summit of the Association of Southeast Asian Nations.

2. A similar result has been observed during the 1997-98 crisis when their competitors' currencies depreciated sharply.

(2002) and de Brouwer (2004) consider the eventuality of an individual basket peg reflecting their own trade structure, before the introduction of a common basket.[3] The aim of a common basket peg would be to reduce the volatility of the nominal effective exchange rate (NEER) in order to preserve Asian countries from changes in their relative competitiveness. Accordingly, a common trade-weighted basket peg would protect the trading relationships among the Asian countries from changes in third-country exchange rates. The proposal to use the ACU -as a coordination mechanism for exchange rate policies- has also gained momentum since the announcement by the Asian Development Bank (ADB) to create a basket of appropriately weighted Asian currencies. For instance, Ogawa and Shimizu (2006a, 2006b) propose the use of an Asian Monetary Unit (AMU) with the aim to monitor Asian exchange rate policies and stabilize their effective exchange rates. By comparing the deviation of each currency vis-à-vis the AMU, they find a misalignment among them and interpret their finding as an illustration of uncoordinated exchange rates policies. Eichengreen (2006) proposes a parallel currency approach with the introduction of an ACU which could play an official role, similar to that played by the European Currency Unit (ECU) within the European Monetary System (EMS).

The coordination of exchange rate policies is crucial for the Asian countries given the level of intra-regional trade and the economic spillovers from potential competitive devaluations and third-currency effects.[4] Indeed, country's authorities might be particularly willing to take into account the movement of neighbor currencies in order to protect their firms from exchange rate's uncertainty and maintain their international competitiveness. This could be achieved through the adoption of a currency basket where the weight of regional currencies would be relatively high. Several studies show that exchange rate volatility may constitute a barrier to trade by increasing currency risk that weighs on firms' profitability and investment decisions.[5] Furthermore, as a result of vertical intra-industry trade in parts, components and semi-finished products, trade structures tend to become similar and the degree of competition among the Asian products tends to increase on third markets (international but also regional markets). Consequently, local firms seek to maintain their market shares by minimizing variation costs and limiting the movement of exchange rates. As argued by Bird and Rajan (2002) and Kawai and Takagi (2005), intra-regional exchange rate stability is therefore necessary to avoid the worsening of terms of trade and promote economic integration in Asia.

In this paper, I examine the extent to which the Asian exchange rates are stabilized against a common basket of regional currencies appropriately weighted by the countries' respective share in the intra-regional trade and the GDP (i.e. the ACU). More specifically, I examine to what extent the movement of the Asian currencies is explained by the movement of the ACU, the US dollar and the euro. By considering the role of the ACU, the analysis conducted in this paper goes beyond the traditional framework of Frankel and Wei (1994) which focus only on major currencies. For this purpose, the econometric tool used for this investigation is a VAR model with Cholesky restrictions, applied to monthly data. Accordingly, one can estimate to what extent each currency is stabilized against other regional currencies and compare the weight of the ACU with those of international currencies (US dollar and euro) in their implicit *de facto* currency basket. This also allows one to take into account a wider range of possibility concerning the authentic currency basket on which the Asian countries peg their currencies.[6] This study

3. For a comparative analysis between individual and common baskets, see Wilson et al. (2007) and Williamson (2009).
4. The intra-regional trade among the ASEAN+3 countries accounts for 49,5% in 2009.
5. Thorbecke (2008) and Chit et al. (2010) find a negative relationship between exports and exchange rate volatility in Asia.
6. Basket peggers generally do not disclose the composition of their currency baskets.

focuses on the nominal exchange rates of South Korea, Indonesia, Malaysia, Singapore, Thailand and the Philippines over the January 2000-March 2011 period.

The results support the hypothesis that the Asian countries stabilize to some extent their exchange rates around the ACU and more specifically after October 2006. Although the US dollar remains the dominant anchor within the region, its decreasing role over these last years leads to conclude that the stability on the US dollar is no longer a priority for these countries. The evidence suggests that the Asian countries have begun a transition process toward a currency basket system where the weight of regional currencies has increased.

The rest of the paper is organized as follows. Section 1.2 reports the methodology to calculate the ACU and presents the econometric model. Section 1.3 presents the estimation results and discussion. Section 1.4 draws conclusions.

1.2 Methodology

Frankel and Wei (1994) popularized a method to identify the weight assigned to major international currencies in the implicit basket peg. Using the Swiss-franc as an independent numeraire, the authors evaluate the extent to which the movements in the Asian exchange rates are explained by the movements in the yen, the mark and the US dollar. The empirical model of Frankel and Wei (1994) is as follows:

$$e_t^{EA} = \alpha_0 + \beta_1 e_t^{USD} + \beta_2 e_t^{EUR} + \beta_3 e_t^{YEN} + \varepsilon_t \tag{1.1}$$

where e is the first difference of the natural logarithm of the respective exchange rates against the Swiss-franc. According to Frankel and Wei (1994), the estimates of β can be interpreted as the respective weights in the implicit basket peg. For instance, if changes in a given currency against the Swiss-franc are mainly explained by the changes in the US dollar against the Swiss-franc, the corresponding coefficient will be close to unity. In this regard, one can conclude that this currency is virtually pegged to the US dollar.

Nonetheless, most of studies using the Frankel-Wei's regression are focusing only on major international currencies, excluding the role of the Asian currencies. Given the reallocation of trade with industrialized countries to intra-zonal trade and the potential spillovers resulting from competitive devaluations and third-currency effects, the Asian countries are likely to directly stabilize their exchange rates against their regional partners and competitors rather than by relying on the US dollar. This could be done through a currency basket where the ACU has a non-negligible weight. By determining the role of the ACU in the management of the Asian exchange rate policies, I check whether this has been the case during the last decade. For this purpose, I introduce the ACU in an extended version of the Frankel-Wei basic model.

Indeed, when introducing the ACU in Eq. (1.1), OLS estimation is biased and inconsistent because the ACU is correlated with the error term. Indeed, the ACU is endogenous as a result of simultaneity with the left-hand side currency because the two variables are co-determined, with each affecting the other. Second, given that the Asian currencies and the ACU are affected simultaneously by the US dollar movements, collinearity arises among the two explanatory variables. Therefore, variance of estimators could be high while the associated t-students could be very low. Moreover, OLS estimators would be highly sensitive to minor changes in the data. Finally, it would be difficult, if not impossible to separate effects of each explanatory variable on the dependent variable.

Therefore, I employ a VAR model with Cholesky restrictions, which represents an appropriate tool to solve endogeneity bias and collinearity issue. I simulate shocks on the external currencies and the ACU to determine the respective share of their innovations (i.e. their implicit weights) in the fluctuation of each Asian currency, by performing variance decomposition and impulse response analyses from the following VAR model:

$$R_t = \phi_0 + \sum_{k=1}^{P} \phi_k(L) R_{t-k} + \varepsilon_t \tag{1.2}$$

where R_t represents the vector of variables ($e^{USD}, e^{EUR}, e^{ACU}, e^{EA_i}$), $\phi_k(L)$ is a (4×4) matrix, and ϕ_0 a vector of constants. Accordingly, the variance decomposition provides the relative weight of each currency (USD, EUR, ACU) in the implicit basket peg of each country (EA_i).

Following Ogawa and Shimizu (2006a), the weight of each currency in the ACU is defined as the arithmetic average of respective countries' share in the GDP (measured at purchasing power parity) and intra-regional trade. These shares are calculated as follows:

$$W_i^{trade} = \frac{X_i + M_i}{\sum(X_i + M_i)} \qquad W_i^{GDP} = \frac{Y_i}{Y_{REG}}$$

with X_i (resp. M_i) the exports from (resp. imports to) country i to (resp. from) other Asian countries, Y_i the GDP of the Asian country i and Y_{REG}, the regional aggregated GDP.[7] These weights are time-varying according to the evolution of the countries' respective share in GDP and intra-regional trade. This is mainly motivated by the rise of China as an important trading partner within the region. The weights are presented in Table 1.1.

Table 1.1: Weights of the Asian currencies in the ACU (in %)

	Indo.	Mal.	Sing.	Thai.	Phil	Viet.	Korea	Japan	China
Periods:									
00-02	5.31	5.97	7.23	5.00	2.66	1.63	11.37	32.67	28.16
03-05	4.91	5.18	7.16	5.00	2.4	1.74	11.24	29.17	33.20
06-08	5.12	4.95	7.27	4.90	2.12	1.99	10.97	25.33	37.37
09-10	5.36	4.72	6.70	4.73	1.95	2.30	10.70	22.95	40.58

Notes: Each row equals to 100%.

The data set cover monthly nominal exchange rates for the January 2000 to March 2011 period ($T = 135$). Following McKinnon and Schnabl (2004), I use low-frequency data because competitiveness could fluctuate sharply from one month to next when the domestic price level is relatively sticky. Furthermore, the incentive to anchor country's price level cannot be recover with high-frequency data because continual changes in exchange rate have little or no effect on domestic prices in the short run. I use the Swiss-franc (CHF) as an independent numeraire to measure exchange rate movements.[8] Bilateral exchange rates are extracted from PACIFIC exchange rate service database.[9]

7. Imports and Exports Data are extracted from the IMF DOTS database and GDP data are extracted from the World Bank database.

8. The exchange rate of the ACU is set at January 2000 = 1 in terms of the US dollar.

9. http://fx.sauder.ubc.ca/

Table 1.2: Test results of the structural changes test in the mean process of the bilateral exchange rates

	idr_usd		myr_usd		php_usd		thb_usd		sgd_usd		krw_usd		cny_usd	
Breaks	BIC	LWZ	BIC	LWZ	BIC	LWZ	BIC	LWZ	BIC	LWZ	BIC	LWZ	BIC	LWZ
0	13.38	13.41	-2.89	-2.85	3.08	3.11	3.08	3.11	-3.66	-3.63	9.77	9.80	-0.88	-0.85
1	13.32	13.38	-4.34	-4.27	2.53	2.60	1.52	1.59	-5.19	-5.12	9.53	9.59	-3.15	-3.08
2	13.27	13.37	-4.60	-4.50	1.99	2.09	1.22	1.32	-5.67	-5.57	8.60	8.70	-4.20	-4.09
3	-	-	-	-	1.93	2.07	-	-	-5.98	-5.84	8.56	8.69	-4.25	-4.12
4	-	-	-	-	-	-	-	-	-6.26	-6.09	-	-	-	-
$SupF_T(1\|0)$	13.93**		456.57***		105.10***		521.69***		505.35***		43.07***		1196.06***	
$SupF_T(2\|1)$	11.37***		44.88***		102.50***		53.82***		89.40***		214.74***		259.09***	
$SupF_T(3\|2)$	-		-		12.86**				53.30***		10.42*		13.05***	
$SupF_T(4\|3)$	-		-		-		-		48.67***		-		-	
					Number of breaks selected									
Sequential	2		2		3		2		4		3		3	
LWZ	2		2		3		2		4		3		3	
BIC	2		2		3		2		4		3		3	
$\hat{T}_1$	2007:11		**2006:11**		2001:08		2003:08		2003:11		2004:10		2004:11	
	(04:06-10:03)		(06:04-06:12)		(01:07-03:10)		(03:01-05-09)		(03:07-04:07)		(04:09-05:05)		(04:08-04:09)	
$\hat{T}_2$	2009:07		2009:07		2003:05		**2006:11**		**2006:03**		**2006:03**		**2006:05**	
	(09:05-13:04)		(07:07-11:07)		(01:08-04:01)		(06:08-07:01)		(06:01-06:05)		(05:04-06:07)		(05:09-06:07)	
$\hat{T}_3$	-		-		**2006:09**		-		2007:09		2008:08		2008:01	
	-		-		(06:05-06:11)		-		(06:12-07:11)		(08:02-08:09)		(07:12-08:04)	
$\hat{T}_4$	-		-		-		-		2009:09		-		-	
	-		-		-		-		(08:10-10:09)		-		-	

Notes: The null hypothesis of $SupF_T(\ell + 1|\ell)$ test is ℓ structural breaks versus the alternative $\ell + 1$ structural breaks. "LWZ" indicates the modified Schwarz criterion of Liu et al. (1997). In parentheses are the 95% confidence interval for the estimated break points. *,**,*** denote significance at 10, 5 and 1 % respectively.

I apply the Bai-Perron methodology (see Bai and Perron , 1998, 2003) for identifying endogenously dates of structural changes in the exchange rate regimes. The structural change analysis is performed on US dollar-based exchange rates rather than CHF-based exchange rates because the latter is assumed to be purely flexible. [10]

I allow up to 4 breaks and use a trimming $\kappa = h/T = 0.15$ with $T = 135$, hence each segment has at least 20 observations ($h = 20$). The results are presented in Table 1.2. For all countries, the estimate detects a break date in 2006 (except for Indonesia), which could be related to the decision of the Chinese authorities to adopt a more flexible regime with reference to an undisclosed basket of currencies. These findings are confirmed by the break test performed on the yuan/dollar exchange rate since we can also observe a break date that take place in 2006. Indeed, the official change of the Chinese exchange rate policy was followed by a yuan's appreciation of 3% during 2006, which is higher than the appreciation observed during the second half of 2005. [11] Since trade with China accounts for an important share of the Asian foreign trade, these countries may have considered such an event in the conduct of their exchange rate policies. [12]. Accordingly, the sample is divided into two sub-samples, one on each side of the 2006 break point (2000:1-2006:9 and 2006:10-2011:3).

Before turning to the VAR analysis, I check for the presence of unit roots in the exchange rate series (in terms of Swiss-franc). The ADF (Augmented Dickey-Fuller) tests indicate that all the variables appear to be integrated of order one, suggesting possible cointegration relationships among them. The results of the Johansen tests indicate no cointegration relationships, so I employ a VAR model in difference as presented in Eq.(1.2) (see the Appendix for ADF and Johansen test results). In the VAR model, the optimal lag length is selected according to the Akaike Information Criteria. Accordingly, the number of lags (p) in the model is one for all countries in the pre-2006 sample and three in the post-2006 sample. The interpretation of shocks is subjected to the identification of structural parameters of the model. Therefore, the Cholesky decomposition is applied to recover the underlying structural shocks by recursive orthogonalization. I constrain the response of the Asian currencies to zero in the face of their respective innovations in order to recover the composition of the currency baskets normalized to one. Finally, I adopt the following causal ordering ($e^{USD}, e^{EUR}, e^{ACU}, e^{EA_i}$) to reflect their level of exogeneity. Here, the assumption is that the US dollar (and the euro) are exogenous to contemporaneous shocks on the ACU.

1.3 Empirical results

1.3.1 Variance decomposition analysis: the role of the ACU

Table 1.3 reports the corresponding forecast error variance decomposition derived from the structural VAR. It shows the corresponding explicative share of structural shocks in the fluctuation of the Asian currencies. The variance decomposition are for 12-month forecast horizon.

10. I consider the case of a pure structural change model. The regression is given by: $y_t = z_t'\delta_j + u_t$ with $t = T_{j-1} + 1, \cdots, T_j$ for $j = 1, \cdots, m + 1$. In this model, y_t is the observed dependent variable at time t; $z_t(q \times 1)$ is the vector of covariates and $\delta_j (j = 1, \cdots, m + 1)$ is the corresponding vector of coefficients; u_t is the disturbance at time t. The indices T_j are the break points. I apply the procedure with only a constant as regressor (i.e. $z_t = 1$) in order to detect structural changes in the mean of the series.

11. Yuan appreciation has begun to accelerate in the mid of 2006 up to October 2008, when the yuan was re-pegged to the US dollar in response to the outbreak of the global financial crisis.

12. For instance, on July 21, 2005, Malaysia quickly followed China and shifted officially from a fixed exchange rate regime to a managed float against an undisclosed basket of currencies.

Table 1.3: Variance decomposition of forecast errors in % of the total variance of the Asian exchange rates.

	2000:01 - 2006:09			2006:10 - 2011:03		
Innovations:	ε_{USD}	ε_{Euro}	ε_{ACU}	ε_{USD}	ε_{Euro}	ε_{ACU}
Malaysia	99.510	0.412	0.078	63.831	7.541	28.628
Indonesia	30.831	31.910	37.259	41.051	28.886	30.062
Singapore	89.806	1.019	9.175	63.492	8.220	28.288
Thailand	78.400	4.341	17.259	63.330	17.474	19.196
The Philippines	90.497	2.293	7.209	60.556	9.9588	29.855
South Korea	66.398	0.351	33.251	27.540	38.797	33.663

Notes: The optimal lag length were selected according to the Akaike Criterion. The lag lengths are 1 and 3 for all countries for the pre- and post-sample periods, respectively.

As a first step, I focus on the estimation results from the first sub-sample. Overall, the explicative share of the ACU is quit low, except for Indonesia and South Korea. Indeed, its explicative share is 0.078%, 37.259%, 9.175%, 17.259%, 7.209% and 33.251% for Malaysia, Indonesia, Singapore, Thailand, the Philippines and South Korea, respectively. Moreover, the US dollar is the dominant anchor in the implicit basket peg of all countries (except for Indonesia). After the currency crisis of 1997-98, it is frequently argued that many Asian countries shifted from rigid currency pegs to managed float systems with varying degree of foreign exchange rate intervention. However, evidence suggests that the these countries have returned to soft US dollar pegging in the aftermath of the 1997-98 crisis. This finding is in line with many empirical studies. For McKinnon and Schnabl (2004), the rationale of the return to official or *de facto* US dollar pegging is its microeconomic role in facilitating international transactions and its macroeconomic role for anchoring regional and national price levels. The return to soft US dollar pegging after the crisis can also result from the need to be competing against neighbors' exporters (i.e. to avoid potential economic spillovers resulting from change in relative prices) who are officially or *de facto* pegged to the US dollar. According to Kenen and Meade (2008), the aversion to exchange rate flexibility derives also from the fear of real appreciation given their export-led growth strategy and competitive pressure in regional and international markets (see, also, Coudert et al. , 2013). In this regard, a common US dollar peg within the region enhances the anchoring effect of any Asian dollar pegger. For Ito et al. (1998) and Ogawa and Ito (2002), this aspect refers to the coordination failure in choosing a desirable exchange rate arrangement since no country would have interest to abandon its US dollar peg as long as other countries continue to stabilize their exchange rates against it.

The estimation results from the second sub-sample display a very different picture. The ACU shocks explain now approximately 30% of the total variances, which is significantly higher than the share in the first sub-sample, especially for Malaysia, Singapore and the Philippines. For other countries, the share of the ACU is stable over the full period and remains relatively high. These results bring evidence that the Asian countries have initiated a shift away from a *de facto* US dollar peg to a currency basket system in which the Asian currencies and the euro have an increasing role. Although the US dollar shocks explain the largest part of the total variance in most cases, the Asian countries have loosened their US dollar pegging since 2006. Indeed, the share of the US dollar has declined for all countries (except Indonesia) in the second sub-sample. For instance, the share of the US dollar has decreased by approximately 40% for Malaysia and South Korea, 30% for Singapore and the Philippines and 15% for Thailand.

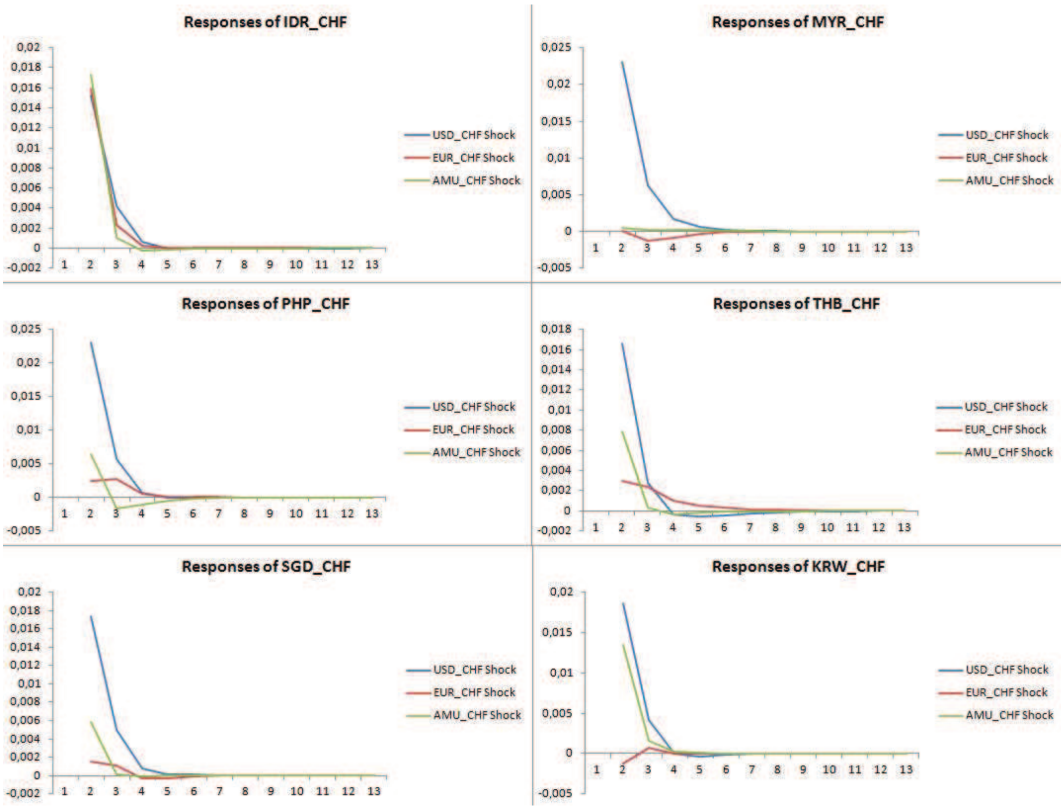

Figure 1.1: Impulse responses of the Asian exchange rates - Period 2000:01-2006:09

The impulse responses to shocks to the US dollar, the euro and the ACU are reported in Figure 1.1 (pre-2006 sample) and Figure 1.2 (post-2006 sample). Figure 1.1 shows that the response to a US dollar shock immediately determines a positive rise in the movement of the home currency. The impulses decrease largely after 2 months and die out after roughly 4-5 months. As expected, the response to the ACU shock is moderate in the first sub-sample (compared to the US dollar shock) for all countries with the exception of Indonesia and South Korea. Furthermore, the response to the euro shock is close to zero for all countries excepted for Indonesia. Concerning the second sub-sample, the magnitude of the response to a shock in the US dollar is smaller for Malaysia, the Philippines, Singapore and South Korea. Finally, the impulse responses to the ACU shock produce an increase in the movement of the exchange rates that becomes negative after roughly 2-3 months. The exchange rates of Malaysia, the Philippines and Singapore are more responsive to innovation in the ACU after 4 months, before finally dying out in the 8th-9th month. This concurs with the variance decomposition results whereby the ACU shocks are larger in determining home currency movements in the second sub-sample.

1.3.2 Is there a yuan effect?

Does the official adoption of a currency basket in China has influenced the other Asian countries? Considering the weight of the yuan in the ACU and the structural breaks observed in 2006, the significant decrease of the US dollar could be attributable to the increasing share of the yuan. This issue might be investigated because the yuan could play a leadership role in the future as a regional monetary

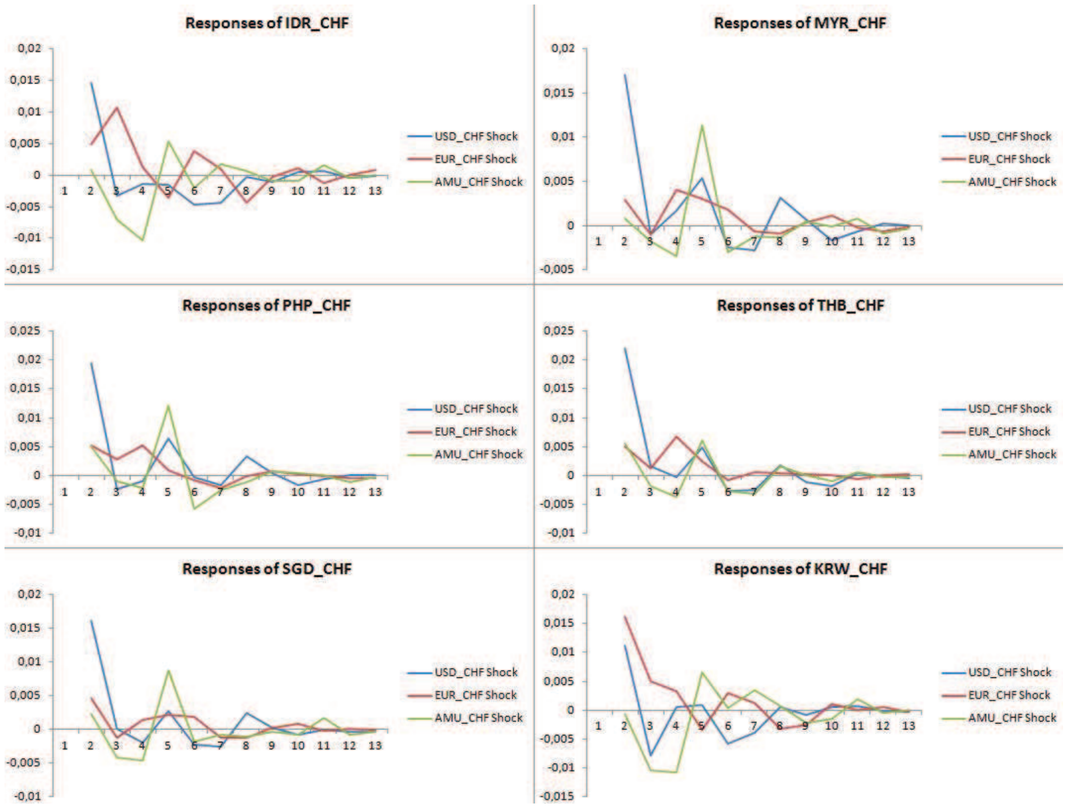

Figure 1.2: Impulse responses of the Asian exchange rates - Period 2006:10-2011:03

anchor.[13] Indeed, the fast pace of the yuan internationalization, along with the China's rise on Asian economic integration, has raised the issue of whether a yuan bloc could be formed within the region. For instance, Park (2010) argues that market integration between China and ASEAN are likely to lead to the emergence of the yuan as an anchor currency. Fratzscher and Mehl (2011) assert that the Chinese exchange rate developments since 2005 are found to exert a strong and growing influence on other Asian exchange rate policies. Accordingly, it would be interesting to analyze the what extent the Asian countries have pegged their currency against the yuan, after China decided to untie its US dollar peg in July 2005. I perform variance decomposition analysis with the yuan instead of the ACU to answer this question. Results are displayed in Table 1.4.

I find some evidence of increasing exchange rate co-movements between the yuan and the Asian currencies since the decision by the Chinese authorities to introduce more exchange rate flexibility. However, it is very difficult to assert that a yuan bloc has emerged in Asia. Indeed, the yuan shocks explain approximately 12% of the total variances in the second sub-sample (except for Indonesia and South Korea where the explicative share of the yuan is 38% and 28%, respectively), which is slightly above compared to the first sub-sample. In other words, the increase in weights of the ACU observed in the preceding section can only be to a certain extent explained by the Chinese currency, thus highlighting the explicative share of other currencies composing the ACU. It would be more appropriate to claim that these countries have allowed for more exchange rate flexibility against the US dollar since 2006, with the aim to adopt a basket peg where the yuan and other Asian currencies have gained an increasing

13. This implies liberalizing and opening its financial system, allowing the yuan's full convertibility and improving the yuan's role in real and financial transactions or foreign exchange reserve holdings.

Table 1.4: Yuan's share in the variance decomposition of forecast errors.

	2000:01 - 2006:09			2006:10 - 2011:03		
Innovations:	ε_{USD}	ε_{Euro}	ε_{Yuan}	ε_{USD}	ε_{Euro}	ε_{Yuan}
Malaysia	99.480	0.412	0.108	79.558	6.566	13.875
Indonesia	43.399	43.879	12.722	41.176	20.045	38.779
Singapore	93.924	1.034	5.042	78.609	10.651	10.74
Thailand	84.478	5.257	10.265	77.233	13.922	8.845
The Philippines	94.148	2.527	3.325	77.142	10.608	12.250
South Korea	76.906	0.346	22.748	29.439	41.876	28.685

Notes: The optimal lag length were selected according to the Akaike Criterion. The lag lengths are 1 and 3 for all countries for the pre- and post-sample periods, respectively. Application of the cointegration test indicates that there is no long-term relationship among the US dollar, the euro, the yuan and the Asian currencies.

role. Given the similarity of their trade-weighted NEER, the Asian countries that peg their currency to a basket are likely to enjoy greater stability across their exchange rates. In this regard, the Chinese exchange rate system reform may have produced greater intra-regional exchange rate stability. This view is also supported by Ma and McCauley (2011) who find that the 2006-2008 experience has rendered the Asian currencies quite stable against each other.

1.4 Concluding remarks

This paper has considered the eventuality of an ACU in the implicit basket peg of several Asian countries to assess the coordination of their exchange rates and recover the composition of their *de facto* basket peg. The key findings of the paper can be summarized as follows. The assessment of the variance decomposition demonstrated that innovations in the US dollar dominate the euro and the ACU shocks after and before the 2006 break date. However, the results also show that the explicative share of the US dollar in the movement of the Asian exchange rates has decreased from roughly 76% to 53% in average, while the explicative share of the ACU has increased from 17% to 29%. Moreover, the decreasing share of the US dollar is also attributable to the euro which has increased from 7% to 18%. These results suggest that an unilateral US dollar peg is no longer a priority for the Asian countries since 2006.

Evidences support the view that the these countries have moved toward a *de facto* currency basket system in which regional currencies play a non-negligible role. Consequently, the recent exchange rate developments in Asia seem to validate many studies which claim that a basket peg would be better suited for them, and that the weight of the US dollar in the aftermath of the 1997-98 crisis was well above its theoretical one (see, e.g., Bird and Rajan , 2002; Bénassy-Quéré , 1999; Ito et al. , 1998). As advocated by Ogawa and Shimizu (2006b), one possible mechanism for strengthening exchange rate coordination inside the region would be to keep a stable relationship with the ACU. This transition step would then pave the way to more advanced forms of monetary integration.

Appendix 1

Table 1.5: Augmented Dickey-Fuller test of stationarity - Period 2000:1- 2006:09

	Intercept	First difference	Intercept and trend	First difference
idr_chf	-1.910	-7.991***	-2.672	-8.022***
krw_chf	-1.821	-6.634***	-1.354	-6.761***
myr_chf	-1.021	-6.717***	-1.977	-6.682***
php_chf	-1.770	-6.670***	-0.433	-6.988***
sgd_chf	-1.321	-7.273***	-0.651	-7.372***
thb_chf	-2.064	-6.923***	-1.441	-7.276***
usd_chf	-0.871	-6.876***	-2.336	-6.824***
eur_chf	-1.689	-7.041***	-2.502	-7.436***
amu_chf	-1.474	-7.279***	-1.063	-7.363***

Notes: *** significant at 1%. The lags were selected through the Schwarz criterion. In all cases the lag is equal to 1. The ADF tests could not reject the null of a unit root in any of these exchange rates in level.

Table 1.6: Augmented Dickey-Fuller test of stationarity - Period 2006:10-2011:03

	Intercept	First difference	Intercept and trend	First difference
idr_chf	-1.465	-6.306***	-1.627	-6.259***
krw_chf	-0.886	-5.798***	-1.301	-5.745***
myr_chf	-1.257	-7.177***	-2.068	-7.127***
php_chf	-0.775	-7.125***	-2.428	-7.168***
sgd_chf	-1.894	-8.341***	-2.854	-8.286***
thb_chf	-1.113	-6.377***	-2.202	-6.406***
usd_chf	-0.452	-6.335***	-1.672	-6.323***
eur_chf	0.923	-6.868***	-1.897	-7.161***
amu_chf	-1.755	-7.148***	-3.648	-7.094***

Notes: *** significant at 1%. The lags were selected through the Schwarz criterion. In all cases the lag is equal to 1. The ADF tests could not reject the null of a unit root in any of these exchange rates in level.

Table 1.7: Cointegration tests - Period 2000:01-2006:09

	Trace Stat.	5% Critical Value	Max-Eigen. Stat.	5% Critical Value
With idr_chf	51,821	63,876	18,160	32,118
With krw_chf	57,762	63,876	29,037	32,118
With myr_chf	53,321	63,876	19,968	32,118
With php_chf	60,515	63,876	24,011	32,118
With sgd_chf	56,583	63,876	27,229	32,118
With thb_chf	52,248	63,876	21,165	32,118

Notes: the other variables are usd_chf eur_chf amu_chf. The tests indicate no cointegration at 5% and the results are robust to lag choice and different deterministic trend specifications.

Table 1.8: Cointegration tests - Period 2006:10-2011:03

	Trace Stat.	5% Critical Value	Max-Eigen. Stat.	5% Critical Value
With idr_chf	52,272	63,876	22,322	32,118
With krw_chf	47,764	63,876	19,507	32,118
With myr_chf	41,064	63,876	17,770	32,118
With php_chf	59,492	63,876	27,875	32,118
With sgd_chf	44,326	63,876	19,065	32,118
With thb_chf	58,632	63,876	27,944	32,118

Notes: the other variables are usd_chf eur_chf amu_chf. The tests indicate no cointegration at 5% and the results are robust to lag choice and different deterministic trend specifications.

CHAPTER

2 Business Cycles Synchronization in East Asia: A Markov-switching Approach

2.1 Introduction

The question of monetary integration in East Asia has become a highly debated issue since the Asian 1997-98 crisis. According to many economists, the peg of their currencies to the US dollar was a source of financial fragility which led to the crisis. As a result, the leaders in the region wondered about the necessity of adopting cooperative financial and monetary policies to protect their economies against a new exchange rate crisis. [1] At the same time, the East Asian authorities agreed on the need to promote a collective arrangement in order to stabilize their exchange rates and foster monetary integration. In the early 2000's governments talked about forming a monetary union with several options that included the constitution of a currency bloc anchored to a common monetary standard. Several options have been suggested and among them a collective pegging to a single currency like the US dollar (McKinnon and Schnabl , 2004), the yen (Kwan , 2001), or the yuan (Park , 2010; Hefeker and Nabor , 2005). Some economists have considered the possibility of a basket peg composed of international currencies (Williamson , 2005) or even regional currencies (Ogawa and Shimizu , 2006a).

From a theoretical point of view, monetary integration in East Asia has been mainly studied in light of the Optimum Currency Areas theory (OCA) which examines the conditions under which two countries can find a gain in adopting a single currency. [2] From the perspective of the OCA theory, the business cycles synchronization and the symmetry of output shocks are crucial because the cost of losing monetary policy independence is considerably reduced when countries experience positively correlated economic shocks. Concerning East Asian countries, the empirical literature is inconclusive on this point. For instance, Bayoumi and Eichengreen (1994) find that supply shocks are symmetric, on the one hand, among Singapore, Malaysia, Indonesia and Hong Kong, and on the other hand, among Japan, Taiwan and South Korea. Chow and Kim (2003) find that domestic outputs of East Asian countries are strongly influenced by country-specific shocks while regional shocks are far more important in European countries that have joined the Economic and Monetary Union. Additionally, Genberg and Siklos (2010) do not clearly identify a group of countries for which shocks are unambiguously highly correlated, while

1. Some initiatives were undertaken with the adoption of reforms aiming, on the one hand at restructuring and increasing the depth of the financial systems (Asian Bond Fund and Asian Bond Market Initiative), and, on the other hand, at setting up mechanisms of protection against speculative attacks (Chiang Mai Initiative).

2. The arguments are based on the comparison of the costs and gains of fixed and flexible exchange rate regimes (see Mundell , 1961, McKinnon , 1963 and Kenen , 1969)

Lee and Koh (2012) suggest that there is a scope among ASEAN countries for potential monetary integration. Also, the authors underline that symmetry of shocks have increased after the 1997-98 financial crisis. Using cointegration techniques, Sato and Zhang (2006) find that some pair-countries in the region share both long-run and short-run synchronous movements of the real outputs, particularly within the ASEAN economies consisting of Singapore, Thailand and Indonesia, and the Northeast Asian region, which consists of Hong Kong, Korea, Mainland China, Japan and Taiwan. Socorro Gochoco-Bautista (2008) analyzes whether the common output fluctuation in East Asia is statistically significant in regressions explaining each country's output fluctuations. The author's findings indicate that there is a significant regional factor that explains the movement of East Asian business cycles. Moneta and Rüffer (2009) find a significant common growth dynamic inside the region by using a state-space framework where common movement is captured by unobservable variables influencing the evolution of the GDP growths. Allegret and Essaadi (2011) introduce a spectral analysis based on the time-varying coherence function and find the presence of a common cycle in East Asia after the crisis of 1997-98, thus suggesting that East Asian countries constitute an OCA. Lee and Azali (2012) examine regional and country-specific cycles simultaneously with the world business cycle by employing the Bayesian state-space Based approach. The authors find an increasing role of region factor although country-specific factor are significant. Finally, some studies have emphasized the role of trade flows in synchronizing economic fluctuations in East Asia (see, e.g., Shin and Sohn , 2006; Rana , 2007; Lee and Azali , 2010).

This paper provides a new approach of business cycle synchronization in East Asia by considering the asymmetric nature of business cycles. We wonder whether cross-country correlations differ between expansions and contractions, or between high-growth and low-growth regimes. We identify which leading cycles are helpful indicators in the prediction of turning points. To this end, we use a Markov-switching model with Time-Varying Transition Probabilities (TVTPMS models). There are two potential causes of business correlation which are important for policy purposes. First, we investigate the hypothesis of recessions and expansions in one country spilling over to other ASEAN-5 countries. The driving forces behind such a phenomenon have been investigated in the literature: trade integration within the ASEAN-5, inter- and intra-industry linkages , co-movements in demand components. These factors explain why the juncture in the ASEAN-5 countries could be driven by a regional joint business cycle. In our framework, this means that an individual country's business cycle can be seen as a leading cycle of the other countries' business cycles. We test this and study the extent to which such a synchronization within the ASEAN-5 area fluctuates over time. However, co-movements between the East Asian countries' business cycles could also reflect the impact of increased globalization, the driving factor being a third country's cycle, typically Japan's, the United States' or China's, since these countries are their main trading partners. We thus examine whether the likelihood to observe a simultaneous switch in the movement of the East Asia's business cycles vary over time according to "international" business cycles. For this purpose, time-varying transition probabilities are also expressed as functions of the GDP growth in the US, Japan and China. Given the leadership role of these countries in driving regional integration process, the extent to which the East Asia's business cycles are affected by the cyclical movements in their economic activity need to be investigated. This issue is particularly important in regard to the possibility of a monetary arrangement with a common currency pegged to the Japanese yen, the Chinese yuan or the US dollar.

The remainder of the paper is organized as follows. Section 2.2 presents the model and describes the data. Section 2.3 contains the empirical results and Section 2.4 concludes.

2.2 Data and methodology

2.2.1 Data

We use quarterly data of the real GDP growth rate of the ASEAN-5 countries, namely Malaysia, Indonesia, the Philippines, Thailand and Singapore. The data begins in the first quarter of 1975 and ends in the second quarter of 2010. The GDP series are taken from the countries' national offices and from the University of Singapore ASU database. We further consider the real GDP growth rates of the US, Japan and China in order to examine the impact of "global" common business cycles. [3]

2.2.2 The model

We want to study whether the activity in the ASEAN-5 countries is driven by a joint business cycle. For this purpose, we proceeds in two steps. We first want to know whether the engine of growth lies within the ASEAN-5 area. We therefore firstly begin by studying the correlations among the ASEAN-5's GDP growth. Then, in a second step, we look for "common factors" referred as the business cycles of their main trading partners (China, Japan and the US).

We chose a measure of the business cycle correlation based on a Markov-switching forewarning model as originally proposed by Filardo (1994), Filardo and Gordon (1998) and reconsidered recently by Kim et al. (2008). This model is useful to answer the following question. Is a country's business cycle, taken as a leading business cycle, informative with regard to detecting the expansion and recession phases of another country's business cycle? The Markov-switching nature of the model is motivated by the fact that we are assuming a notion of causality in the GRANGER sense (not only correlation) between business cycles: the business cycle A causes the business cycle B, if the information in A helps predicting B. A is the leading cycle. Assume that A refers to the phases of expansion or recession observed in a country at a given time *t* and that we want to know their influence on B *t+s* period ahead. Since the expansion and recession regimes in B are not yet observed when we are in time *t*, they are considered as hidden states to which we assign a given probability of realization.

The basic model : a regional cycle within the ASEAN-5 area

In the first formulation of the TVTPMS model, we study the co-movements among the business cycles of the ASEAN-5. We consider the effects of expansions and recessions in one country spilling over to the other countries.

Let y_t be the growth rate in a given country at time t and z_t the growth rate in another country at time t. We want to know whether z_t causes y_{t+k} , $k = 0, 1, 2, ...$. Under the assumption that both y and z have ergodic distributions, we can adopt a "backward" reasoning to investigate whether y at time t is the result of the dynamics of z during the preceding k periods. The model has the following parametric form:

3. All GDP data have been seasonally adjusted. The GDP series for China begins in 1978:2.

$$y_t = \mu_1 + \sum_{m=1}^{M} \phi_m y_{t-m} + \sigma\epsilon_t \quad \text{in regime 1} \tag{2.1}$$

$$= \mu_2 + \sum_{m=1}^{M} \phi_m y_{t-m} + \sigma\epsilon_t \quad \text{in regime 2} \tag{2.2}$$

μ_1, μ_2, σ, ϕ_m are real coefficients to be estimated. ϵ is distributed as a Normal law $N(0, 1)$. We define s_t as an unobservable or hidden variable governed by a two-regime Markov-chain of order 1 with the following time-varying transition probability matrix

$$P_{ij}(s_t = i|s_{t-1} = j, z_{t-k}) = \begin{pmatrix} P_{11}(z_{t-k}) & 1 - P_{22}(z_{t-k}) \\ 1 - P_{11}(z_{t-k}) & P_{22}(z_{t-k}) \end{pmatrix}, \tag{2.3}$$

with P_{ij} the probability of switching from regime j at time $t-1$ to regime i at time t and $i, j = 1, 2$ with $\sum_{j=1}^{2} P_{ij} = 1$ for all $i, j \in \{1, 2\}$. k is a lag. The functional form linking z_{t-k} to P_{ij} is logistic:

$$P_{11}(z_{t-k}) = \frac{\exp(\theta_{1,1} + \theta_{1,2}z_{t-k})}{1 + \exp(\theta_{1,1} + \theta_{1,2}z_{t-k})} \quad \text{and} \quad P_{22}(z_{t-k}) = \frac{\exp(\theta_{2,1} + \theta_{2,2}z_{t-k})}{1 + \exp(\theta_{2,1} + \theta_{2,2}z_{t-k})}, \tag{2.4}$$

These probabilities provide information about the likelihood of staying or switching from a given regime (either regime 1 or regime 2) k periods after a change has occurred in z. Since y and z are growth rates, regimes 1 and 2 capture expansions and recessions, or low- and high-growth rate regimes. The latter are not selected *a-priori*, but determined endogenously by the data. Let us consider an example. Suppose that the data yields an estimated coefficient μ_1 which is a positive and an estimate of μ_2 which is negative. This would indicate that regime 1 can be interpreted as one of expansion and regime 2 as one of recession. Assume further that in Eq. (2.4) $\theta_{1,2}$ is positive. This means that any increase (resp. decrease) in z increases P_{11}, i.e. the probability that y stays in regime 1 (resp. $1 - P_{11}$, i.e the probability that y switches from regime 1) k periods later: an expansion (resp. recession) in a leading country yields an expansion (resp. recession) in another country. A negative coefficient would indicate that an expansion in a leading country causes a recession in another country or reduces the likelihood that the country will continue to evolve in a regime of expansion. We have a similar interpretation for $\theta_{2,2}$. For instance a negative coefficient would indicate that any decrease (resp. increase) in z increases the probability of staying in regime 2 (resp. switching from regime 2). It can happen that both coefficients $\theta_{1,2}$ and $\theta_{2,2}$ are insignificant. This would mean that the business cycle in a leading country is uninformative of the occurrence of expansions and recessions in another country. This assumption is also tested using a log-likelihood ratio test with the following statistic

$$LR = 2 \times [L_{TVTP}(\Theta) - L_{FTP}(\Theta)]$$

where $L_{TVTP}(\Theta)$ and $L_{FTP}(\Theta)$ are respectively the log-likelihood of the TVTPMS model and the fixed transition probabilities (FTPMS) model of Hamilton (1989). If the null hypothesis of fixed transition probabilities is not rejected at conventional significance levels, the TVTPMS model converges to the FTPMS model. In such a case, the information variable does not help predict future changes in regimes.

The extended model

In the basic model, we assumed that the business cycle in a country is driven by a "regional" business cycle. This means that in the basic model the countries considered to examine the business cycle correlation are members of the ASEAN-5. But co-movements in their business cycles can also result from a common cycle outside their region. We therefore consider the following extended model:

$$y_t = \mu_1 + \sum_{m=1}^{M} \phi_m y_{t-m} + \beta_1 x_t + \sigma\epsilon_t \quad \text{in regime 1} \tag{2.5}$$

$$= \mu_2 + \sum_{m=1}^{M} \phi_m y_{t-m} + \beta_2 x_t + \sigma\epsilon_t \quad \text{in regime 2} \tag{2.6}$$

The transition probabilities are defined in a similar way as in the basic model. In Eq. (2.6) y and x refers to the GDP growth rate of two ASEAN-5 countries and their degree of synchronization are captured by the coefficients β_1 and β_2. The signs of the constants μ_1 and μ_2 still indicate whether the observed regimes correspond to expansions and recessions, or high and low-growth regimes. The variable z refers to the GDP growth rate of China, Japan or the US.

Estimation

We present the estimation procedure for the extended model since the latter encompasses the basic specification. We employ the maximum likelihood method to provide estimates of the parameters [4]. We define the vector of all the parameters in the model as θ , $\xi_t = (y_t, y_{t-1}, \cdots, y_1)'$, the complete history of the endogenous variable y observed through date t and $\Omega_t = (X'_t, Z'_t)'$, the vector observations of the exogenous variable x and the information variable z through date t. The conditional log-likelihood function of the observed data is given by the log-density of y_t given ξ_{t-1} and Ω_t,

$$L(\theta) = \sum_{t=1}^{T} \ln f(y_t|\Omega_t, \xi_{t-1}; \theta) \tag{2.7}$$

obtained using the following joint density-distribution,

$$f(y_t|\Omega_t, \xi_{t-1}; \theta) = \sum_i \sum_j f(y_t, s_t = i, s_{t-1} = j|\Omega_t, \xi_{t-1}; \theta) \tag{2.8}$$

$$= \sum_i \sum_j f(y_t|s_t = i, s_{t-1} = j, \Omega_t, \xi_{t-1}; \theta) \cdot P(s_t = i, s_{t-1} = j|\Omega_t, \xi_{t-1}; \theta) \tag{2.9}$$

We also have

$$P(s_t = i|\Omega_t, \xi_t) = \frac{\sum_j f(y_t = i, s_{t-1} = j|\Omega_t, \xi_{t-1}; \theta)}{f(y_t|\Omega_t, \xi_{t-1}; \theta)} \tag{2.10}$$

4. The BFGS algorithm is used to perform nonlinear optimization

that describes how time variation in the transition probabilities affects the probability of the unobserved regimes of the economy at time t (probability of being in either regime 1 or regime 2). Finally we need the regime-dependent conditional density function. Under the assumption that ϵ_t is distributed as a Normal law, the density of y_t conditional to s_t is

$$f(y_t|s_t = 1, s_{t-1} = j, \Omega_t, \xi_{t-1}; \theta) = \frac{1}{\sigma\sqrt{2\pi}} \exp\left\{-\frac{1}{2}\left(\frac{(y_t - \hat{y}_t)}{\sigma}\right)^2\right\} \tag{2.11}$$

$$f(y_t|s_t = 2, s_{t-1} = j, \Omega_t, \xi_{t-1}; \theta) = \frac{1}{\sigma\sqrt{2\pi}} \exp\left\{-\frac{1}{2}\left(\frac{(y_t - \hat{y}_t)}{\sigma}\right)^2\right\} \tag{2.12}$$

with

$$(y_t - \hat{y}_t) = \frac{y_t - \mu^{s_t} - \beta^{s_t} x_t - \phi_m(y_{t-m} - \mu^{s_{t-1}} - \beta^{s_{t-1}} x_t)}{\sigma} \tag{2.13}$$

Finally, the lags in the model are chosen using the Akaike information criterion. Furthermore, we perform the Ljung-Box (LB) test to check that is no residual autocorrelation. In all cases, the Q-statistics indicate that residuals are independently distributed. [5]

2.3 Empirical Results

2.3.1 Preliminary evidence of business cycle synchronization

Figures 2.1 till 2.5 provide some preliminary evidence of business cycle commonalities in the ASEAN-5 (see Appendix 1). The figures display the smoothed probabilities P_{22} estimated from a standard Hamilton (1989) model. We note that, in general, the computed probabilities overlap periods usually referred as times of contractions in the GDP growth especially during the years 1984-1985, 1997-1998 and 2008-2009. Besides, the five countries share episodes of continuous growth expansion over the years 1986-1995 and the mid-2000s. The results suggest that business cycles in the ASEAN-5 are well-tied together. Table 2.1 reports the dating of regimes based on the smoothed probabilities. The average duration of expansions is approximately five years for Singapore and Malaysia, four years for Indonesia and three years for Thailand and the Philippines. Conversely, the average duration of contractions is less than one year for all the countries except Malaysia. This reveals some evidence of a symmetry between the expansion and contraction phases.

2.3.2 Results from the basic TVTPMS model

Tables 2.2 till 2.6 report the estimates of the TVTPMS basic model by country. We find significantly positive μ_1 and negative μ_2 which correspond to a situation of distinct expansion and recession regimes. Regarding the values of the coefficients, the positive coefficients captures increases in the GDP growth from 1,1 per cent per quarter (in the Philippines) to 2,2 per cent (in Singapore) on average, which amounts to annual growth rates ranging on average from 4,4 per cent to 9 per cent. Indonesia seems to

5. Results are reported in the next section.

Table 2.1: Duration of expansions and recessions

	Growth-expansion regime		Growth-contraction regime	
Singapore	1977:02-1985:01	2001:04-2003:01	1985:02-1985:04	2003:02
	1986:01-1997:03	2003:03-2008:01	1997:04-1998:03	2008:02-2009:01
	1998:04-2000:04	2009:02-2010:02	2001:01-2001-03	
Average duration	19,6 quarters		3 quarters	
Thailand	1977:02-1979:02	1995:04-1996:03	1979:03	1996:04-1997:01
	1979:04-1980:01	1997:02	1980:02	1997:03-1998:03
	1980:03-1985:02	1998:04-2008:02	1985:03	2008:03-2009:01
	1985:04-1991:03	2009:02-2010:01	1991:04	2010:02
	1992:01-1995:02		1995:03	
Average duration	13 quarters		1,77 quarters	
The Philippines	1977:02-1979:03	1992:03-1997:04	1977:02	1991:01-1991:02
	1980:02-1981:01	1999:01-2000:03	1979:04-1980:01	1992:02
	1981:03-1983:01	2001:04-2008:02	1981:02	1998:01-1998:04
	1986:01-1990:04	2009:02-2010:02	1983:02-1985:04	2000:04-2001:03
	1991:03-1992:01		1986:04	2008:03-2009:01
Average duration	12,1 quarters		2,66 quarters	
Malaysia	1978:03-1984:02	2001:04-2008:02	1977:02-1978:02	2000:04-2001:03
	1986:02-2000:03	2009:01-2010:02	1984:03-1986:01	2008:03-2008:04
Average duration	22 quarters		4,6 quarters	
Indonesia	1977:02-1982:03	1995:01-1996:03	1982:04-1983:01	1994:01-1994:04
	1983:02-1984:04	1997:02	1985:01-1985:02	1996:04-1997:01
	1985:03-1992:01	1999:02-2010:02	1992:02-1992:04	1997:03-1999:01
	1993:01-1993:04			
Average duration	16,1 quarters		3,33 quarters	

experience recessions of a larger magnitude (8 percent on average per quarter) than the other countries (the smallest amplitude is observed for the Philippines with an average of 1,3 per cent per quarter). The evidence of an asymmetry among the magnitude of expansions and recessions can be explained by the fact that the latter incorporates the years corresponding to the 1997-1998 and 2008-2009 crises.

For purpose of illustration, we proceed to comment upon the estimated results when the Malaysian business cycle is considered as the leading cycle (the variable z_{t-k}). The coefficient $\theta_{1,2}$ is insignificant in the case of Thailand and Indonesia, thereby implying that a higher growth rate in Malaysia is not informative of the changes observed in these countries' GDP growth during expansions. For the Philippines and Singapore this coefficient is significantly positive, which is indicative of the fact that an expansion in Malaysia increases the probability that these countries will evolve in an expansion regime (i.e P_{11}). Alternatively, it means that a recession in Malaysia increases the probability that these countries will fall into the contraction regime (i.e. $1 - P_{11}$). We see that $\theta_{2,2}$ is never significant. This suggests that the Malaysian cycle can never be considered as a leading indicator of the future state of the cycle in the other ASEAN-5 countries when they are already in the contraction regime. Figure 2.6 plots the value of the time-varying probabilities P_{11} of Thailand, Indonesia, the Philippines and Singapore implied by changes in the GDP growth of Malaysia (see Appendix 1). The results show that Malaysia is not predictive of the probability of staying in the growth-expansion regime of Thailand and Indonesia because the probability P_{11} remains fairly high regardless of innovations in z_{t-k}. This is not a surprising result when considering that $\theta_{1,2}$ is not statistically significant for these countries. For the situation of other countries, we see that a decrease in the GDP growth decreases the probability P_{11}, thereby suggesting that a shift from regime 1 is more likely to occur following a contraction of economic activity in Malaysia. When examining the slop of the logistic trend functions, it seems that Singapore is less sensitive to negative changes than the Philippines.

The interpretation of the coefficients $\theta_{1,2}$ and $\theta_{2,2}$ is similar when the other countries are considered as the leading countries (in terms of business cycle fluctuations). Taken as a whole, the information conveyed by the different ASEAN-5 growth variables for the expansion and recession phases in a given country is the following. For each country, we look at the significance of the parameters $\theta_{1,2}$ and $\theta_{2,2}$. There are two striking features from the estimates. First, spillover business cycle effects in the ASEAN-5 explain little the expansion phases in Indonesia, since for this country the coefficients $\theta_{1,2}$ are insignificant when the other countries' GDP growths are considered as the leading indicators. However, its business cycle contribute in general to the observed changes during expansions in the other countries. For Indonesia these observations suggest an unidirectional causality from its business cycle to the others.

Table 2.2: Estimations results for Malaysia as endogenous and regional cycles

z Lag k	Indonesia $t-0$	The Philippines $t-2$	Thailand $t-0$	Singapore $t-1$
μ_1	1,821***	1,809***	1,807***	1,811***
	(14,41)	(15,27)	(16,20)	(15,87)
μ_2	-2,721***	-2,972***	-2,992***	-2,884***
	(-5,94)	(-6,46)	(-6,16)	(-5,43)
ϕ	0,10	0,06	0,06	0,05
	(1,28)	(0,72)	(0,75)	(0,58)
σ	1,247***	1,237***	1,236***	1,243***
	(18,58)	(17,64)	(18,00)	(16,87)
$\theta_{1,1}$	3,28***	3,844***	4,095***	2,976***
	(4,76)	(5,76)	(3,94)	(4,95)
$\theta_{2,1}$	1,92	0,54	1,69	0,48
	(0,91)	(0,66)	(1,46)	(0,44)
$\theta_{1,2}$	**0,839***	**0,56****	**0,987*****	**0,973****
	(1,91)	(2,44)	(3,09)	(2,43)
$\theta_{2,2}$	-1,33	0,88	**-1,626****	-0,01
	(-0,95)	(1,22)	(-2,09)	(-0,01)
LR	11,23	6,45	20,40	7,00
	[0,00]	[0,04]	[0,00]	[0,03]
$Q(1)$	0,10	0,25	0,22	0,17
	[0,75]	[0,62]	[0,64]	[0,68]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.]

Table 2.3: Estimations results for the Philippines as endogenous and regional cycles

z Lag k	Malaysia $t-1$	Indonesia $t-0$	Thailand $t-2$	Singapore $t-0$
μ_1	1,119***	1,135***	1,135***	1,072***
	(8,46)	(8,83)	(8,76)	(7,43)
μ_2	-1,753***	-1,109***	-1,339***	-1,323**
	(-3,51)	(-2,87)	(-3,48)	(-2,49)
ϕ	-0,19**	-0,179***	-0,216**	-0,167**
	(-2,18)	(-2,90)	(-2,61)	(-2,48)
σ	1,67***	1,696***	1,683***	1,729***
	(15,75)	(16,77)	(16,15)	(16,63)
$\theta_{1,1}$	2,879***	5,912**	3,59***	4,95***
	(3,06)	(1,99)	(3,47)	(2,90)
$\theta_{2,1}$	0,10	3,224*	2,43	1,784*
	(0,08)	(1,85)	(1,38)	(1,69)
$\theta_{1,2}$	**1,438****	**3,652***	**0,987****	-0,19
	(2,30)	(1,66)	(2,27)	(-0,32)
$\theta_{2,2}$	1,10	-0,73	7,14	0,21
	(0,76)	(-1,30)	(0,84)	(0,24)
LR	6,66	8,35	7,31	0,15
	[0,04]	[0,02]	[0,03]	[0,93]
$Q(1)$	1,10	0,28	0,23	0,43
	[0,31]	[0,60]	[0,63]	[0,51]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.4: Estimations results for Singapore as endogenous and regional cycles

z Lag k	Malaysia $t-1$	Indonesia $t-0$	The Philippines $t-1$	Thailand $t-1$
μ_1	2,194***	2,184***	2,215***	2,195***
	(17,39)	(17,02)	(17,19)	(17,48)
μ_2	-1,561***	-1,512***	-1,274***	-1,482***
	(-3,94)	(-3,38)	(-3,27)	(-3,47)
ϕ	0,03	0,01	-0,02	0,02
	(0,25)	(0,05)	(-0,18)	(0,16)
σ	1,344***	1,364***	1,365***	1,355***
	(15,51)	(15,45)	(16,03)	(15,41)
$\theta_{1,1}$	2,048***	2,507***	3,178***	2,611***
	(3,55)	(4,05)	(5,15)	(4,73)
$\theta_{2,1}$	0,50	0,79	0,99	0,52
	(0,87)	(1,28)	(1,62)	(0,78)
$\theta_{1,2}$	**0,908****	**0,813***	**0,5****	0,45
	(2,31)	(1,82)	(1,98)	(1,38)
$\theta_{2,2}$	-0,18	-0,20	0,12	-0,42
	(-0,67)	(-0,65)	(0,25)	(-1,15)
LR	5,46	5,39	2,54	3,85
	[0,07]	[0,07]	[0,28]	[0,15]
$Q(1)$	0,001	0,03	0,002	0,001
	[0,98]	[0,85]	[0,96]	[0,98]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.5: Estimations results for Indonesia as endogenous and regional cycles

z	The Philippines	Thailand	Singapore	Malaysia
Lag k	$t-1$	$t-2$	$t-2$	$t-1$
μ_1	1,454***	1,454***	1,454***	1,45***
	(15,76)	(15,94)	(15,77)	(16,04)
μ_2	-7,927***	-7,927***	-7,927***	-7,92***
	(-9,74)	(-9,83)	(-9,79)	(-11,17)
ϕ	-0,11	-0,11	-0,112*	-0,108*
	(-1,30)	(-1,31)	(-1,72)	(-1,92)
σ	1,197***	1,197***	1,197***	1,189***
	(16,57)	(16,64)	(20,25)	(20,66)
$\theta_{1,1}$	5,269***	4,684***	5,236***	4,701***
	(3,88)	(4,48)	(3,25)	(6,92)
$\theta_{2,1}$	3,05	2,17	3,42	-65,75
	(0,88)	(0,56)	(1,28)	(-0,01)
$\theta_{1,2}$	-0,20	0,51	-0,10	0,15
	(-0,37)	(1,21)	(-0,15)	(1,44)
$\theta_{2,2}$	2,41	0,51	2,36	-14,56
	(0,86)	(0,56)	(1,21)	(-0,01)
LR	0,55	1,40	1,76	2,84
	[0,76]	[0,50]	[0,41]	[0,24]
$Q(1)$	1,06	1,06	1,06	0,99
	[0,30]	[0,30]	[0,30]	[0,32]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.6: Estimations results for Thailand as endogenous and regional cycles

z	Malaysia	Indonesia	The Philippines	Singapore
Lag k	$t-1$	$t-0$	$t-1$	$t-1$
μ_1	1,713***	1,683***	1,712***	1,771***
	(12,16)	(14,32)	(12,22)	(12,61)
μ_2	-3,428***	-4,062***	-3,456***	-2,908***
	(-3,52)	(-5,52)	(-3,70)	(-4,40)
ϕ	0,02	-0,02	0,02	-0,05
	(0,18)	(-0,25)	(0,21)	(-0,43)
σ	1,449***	1,457***	1,449***	1,42***
	(15,16)	(17,29)	(15,15)	(13,90)
$\theta_{1,1}$	3,775***	2,839***	3,545***	2,592***
	(3,68)	(5,64)	(5,58)	(3,97)
$\theta_{2,1}$	-0,08	-54,98	0,10	-0,13
	(-0,06)	(0,00)	(0,11)	(-0,16)
$\theta_{1,2}$	-0,09	**0,849*** **	0,12	**0,678** **
	(-0,18)	(19,73)	(0,37)	(1,98)
$\theta_{2,2}$	-0,06	-12,40	-0,15	-0,29
	(-0,17)	(0,00)	(-0,39)	(-0,68)
LR	0,07	6,15	0,32	3,04
	[0,97]	[0,05]	[0,85]	[0,22]
$Q(1)$	0,02	0,06	0,03	1,03
	[0,88]	[0,80]	[0,86]	[0,31]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

The second striking results is that, in general, recessions in any country have no discernible effects on other countries when they are already into recession. Indeed, most of the time, the coefficient $\theta_{2,2}$ is insignificant, thereby implying that any change in a given GDP growth does not help predict whether other economies will stay into or escape from recessions. This second finding is consistent with two potential explanations. While the increasing economic interdependence -through intra-regional trade and direct investments- tends to strengthen output co-movements when the countries are already in the expansionary state, the shift from recessions to recovery are mainly due to idiosyncratic policies (e.g. monetary, fiscal and wage policies). Nevertheless, this does not exclude the possibility for countries' cycle to be correlated during turbulent periods as demonstrated in 1997-98. This finding is supported by some studies which claim that the East Asian countries still need cooperative mechanisms to strengthen their ability to manage crises, and insulate countries from policy spillovers such as "beggar-thy-neighbor" competitive devaluations (see, e.g., Plummer , 2006 and Volz , 2006). An alternative explanation may be that there do exist a correlation among the contraction phases of the business cycles caused by a global factor that is not accounted for here. We now look at this second explanation.

2.3.3 Results from the extended model

Tables 2.8 till 2.12 report the estimation for the extended model (see Appendix 2). The main difference with the basic model is that the leading cycles are now those of Japan, China and the US, and, the ASEAN-5 growth rates enter as exogenous variables in each country regression. Studies on macroeconomic linkages among the East Asian countries highlight that East Asian business cycles have become more responsive to the economic cyclicality of the developed countries where Asian exports are mainly consumed. Indeed, global economic linkages among the East Asian countries and the rest of the world are especially explained by the heavy dependence on final demand in industrial economies, especially with China, the US and Japan (see Table 2.7 in Appendix 2).

The TVTP model allows taking into consideration how the economic cyclicality of a third country affects business cycles synchronization in the ASEAN-5. The model still dichotomizes between two regimes of expansion and recession, since we find statistically significant positive and negative values for the coefficients μ_1 and μ_2. The coefficients β_1 and β_2 capture the correlation among the ASEAN-5's business cycles and the latter are assumed to be different during recessions and expansions. It is found that the correlations among the ASEAN-5's cycles are stronger during contractions than during expansions (the estimated coefficients β_2 are much larger than β_1). Therefore, a drop in the growth rate of Japan, the US and China, could explain that the ASEAN-5 countries fall into recession a few quarters later. This would be suggestive of a global international cycle driving recessions in these countries.

Regarding the influence of the Japanese, Chinese and American cycles on the ASEAN-5 business, we summarize in Table 2.13 (see Appendix 2) our main findings based on the significance of the estimated coefficients $\theta_{1,2}$ and $\theta_{2,2}$. The results shows a great heterogeneity of situations. In some cases, the "international" cycles do not provide significant content in explaining the ASEAN-5 business cycle contractions and expansions, because some useful information about the dynamics of the growth rate is already provided by another ASEAN-5 country's business cycle. This suggests a dominance of a regional cycle over a global cycle. For instance, this happens in Malaysia. Considering the regression in which Singapore's growth rate enters as an explanatory variable, we find significant coefficients β_1 and β_2 and an insignificant coefficient $\theta_{1,2}$ when the Japanese' cycle is chosen as the leading international business cycle. But, this coefficient become statistically significant when the leading cycle is the Chinese's and the American's.

In most regressions, the Indonesian cycle captures only part of the influence of the external growth on the ASEAN-5 countries' growth rates when one also considers the role of the American and Chinese cycles. Indeed, the parameter $\theta_{1,2}$ are significant in the regressions. However, the Indonesian cycle captures the influence of the Japanese business cycle on the growth rates of the ASEAN-5 countries (in this case, the coefficient is insignificant). We see that the business cycle in Thailand captures most of the influence of the "international" cycles, since in a majority of regressions in which the growth rate of this country enters as an explanatory variables, the parameter $\theta_{1,2}$ is insignificant. A similar argument applies for the Philippines. These observations play in favor of the view that the business cycles among the ASEAN-5 are correlated through a regional common cycles, rather through an influence of a common "international cycle". Accordingly, the synchronization process is enhanced by strong intra-regional linkages (e.g. through capital, intermediate goods, services and labor markets) that transmit macroeconomic fluctuations among interdependent partners, from larger economies to smaller one. This view is consistent with the findings that the synchronization process is mainly driven by major economies -in terms of PPP-based GDP- in Southeast Asia, that is Indonesia and to a lesser extent Malaysia and Thailand. Moreover, changes in the Chinese, Japanese and American cycles are uninformative of the likelihood to evolve in or escaping from a regime of contraction in the countries since the parameter $\theta_{1,2}$ is most often insignificant. The fact that "international cycle" are uninformative during contraction episodes leads to conclude that the adjustment process in response to shocks display idiosyncratic features as suggested by our preceding results.

2.4 Conclusion

The analysis in this paper suggests that a regional cycle of the ASEAN-5 could provide significant informational content in predicting the future state of the economies only when they are already into the expansionary state. Thus the causality between the business cycles is asymmetric. Accordingly, an ASEAN-5 monetary union appears to be premature since the Southeast Asian countries are not fully synchronized with each other despite evidence of business cycle transmission and interdependence within the region. This calls for a deeper regional economic cooperation, including intra-exchange rate stability and macroeconomic policy coordination, before turning on to a full-fledged monetary union. Furthermore, our finding plays against the idea according to which the ASEAN-5 countries' economies are characterized by a strong dependence on external demand. The estimates appear to suggest that they instead share a common region-specific business cycle, with a driving role for Indonesia, the country with the highest GDP within the region. The Asian regional trade may have played an increasingly role in strengthening the business correlation within the group. This provides support to the arguments that the ASEAN-5 could decouple from the biggest countries and manage to generate a self-sustaining growth path. There are several directions in which this study could be extended. First, it would be interesting repeating a similar exercise with the components of the GDP growth (trade, investment, consumption) and with monetary and financial variables in order to investigate other sources of correlation between the countries. Secondly, our model focus on cross-country spillover effects within the region. It would be interesting looking at the effects of common exogenous shocks by adding features on impulse response functions from Markov-switching models.

Appendix 1

Figure 2.1: Smoothed probabilities of being in a regime of contraction: Singapore. Upper panel: GDP growth rate, lower panel: smoothed probabilities

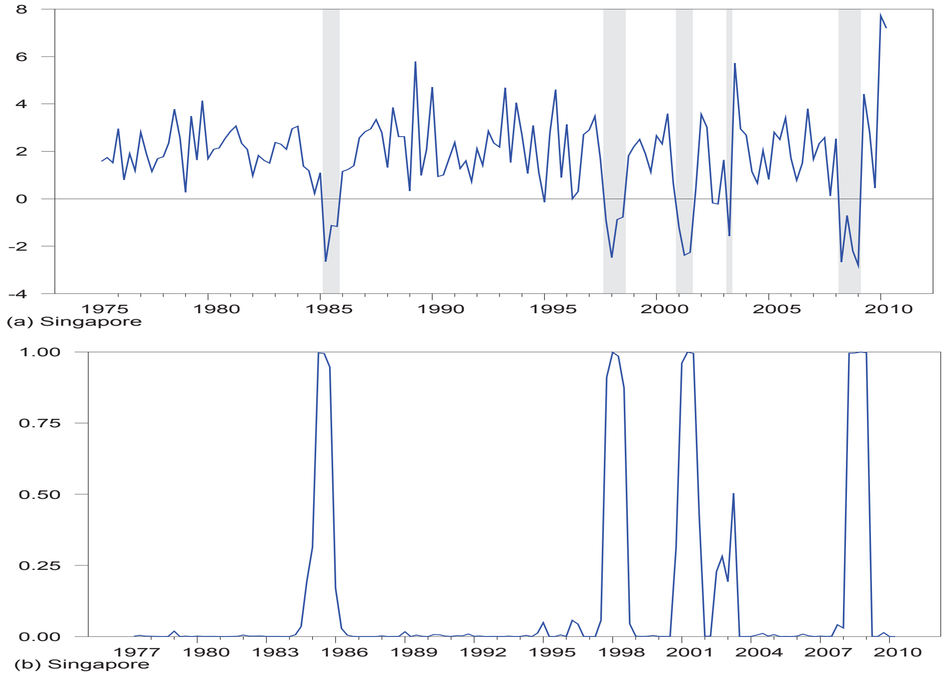

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Figure 2.2: Smoothed probabilities of being in a regime of contraction: Malaysia. Upper panel: GDP growth rate, lower panel: smoothed probabilities

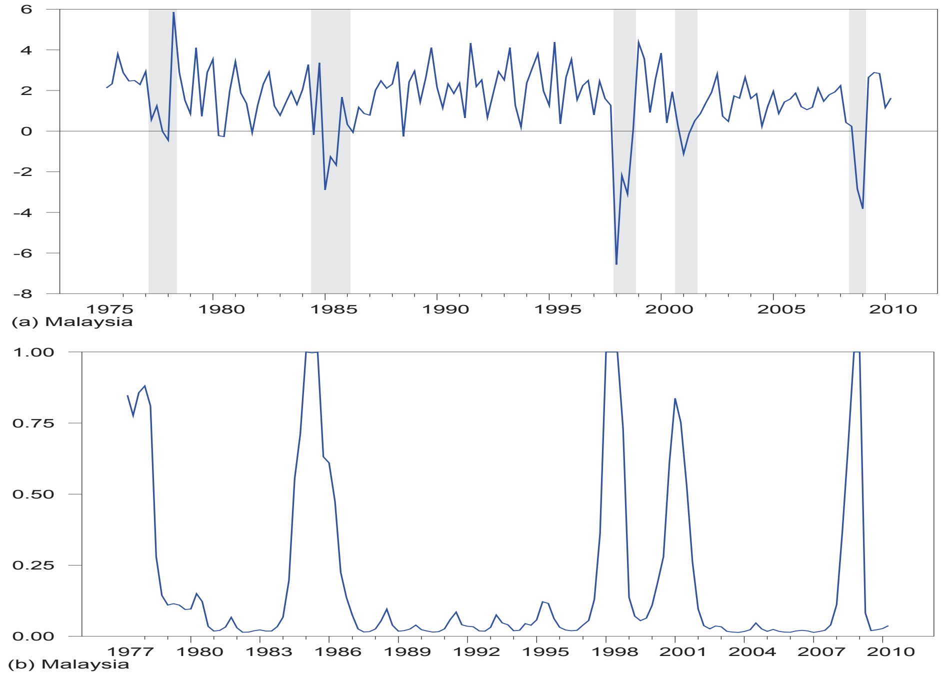


Figure 2.3: Smoothed probabilities of being in a regime of contraction: the Philippines. Upper panel: GDP growth rate, lower panel: smoothed probabilities

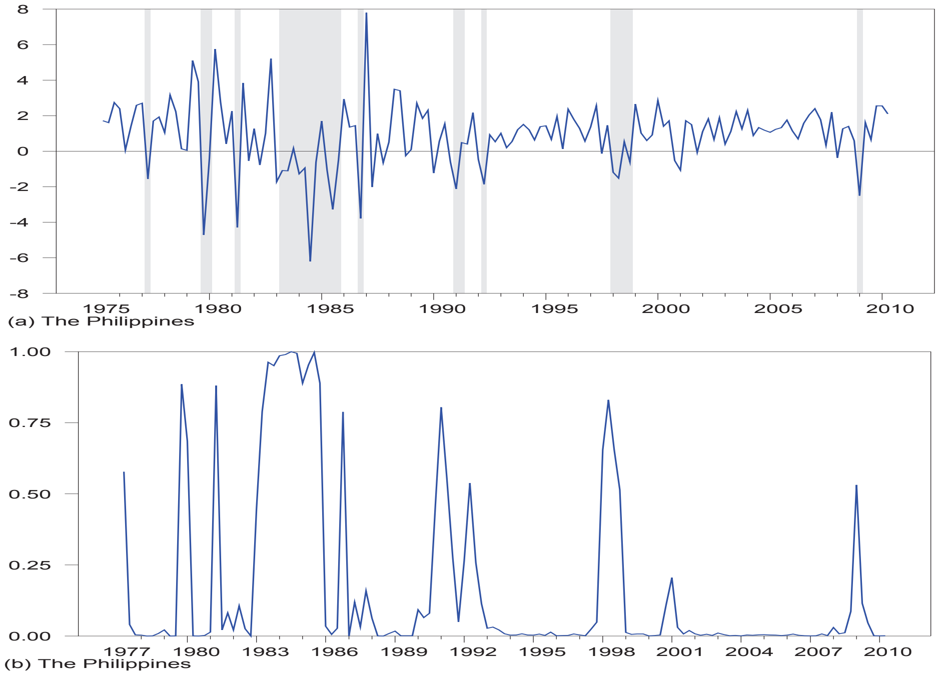


Figure 2.4: Smoothed probabilities of being in a regime of contraction: Thailand. Upper panel: GDP growth rate, lower panel: smoothed probabilities

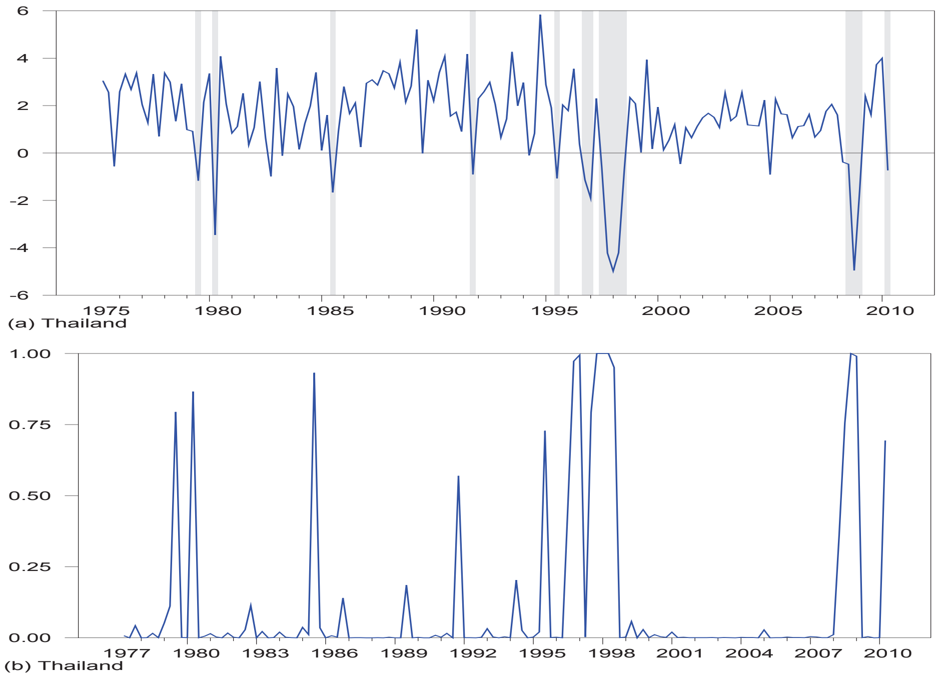


Figure 2.5: Smoothed probabilities of being in a regime of contraction: Indonesia. Upper panel: GDP growth rate, lower panel: smoothed probabilities

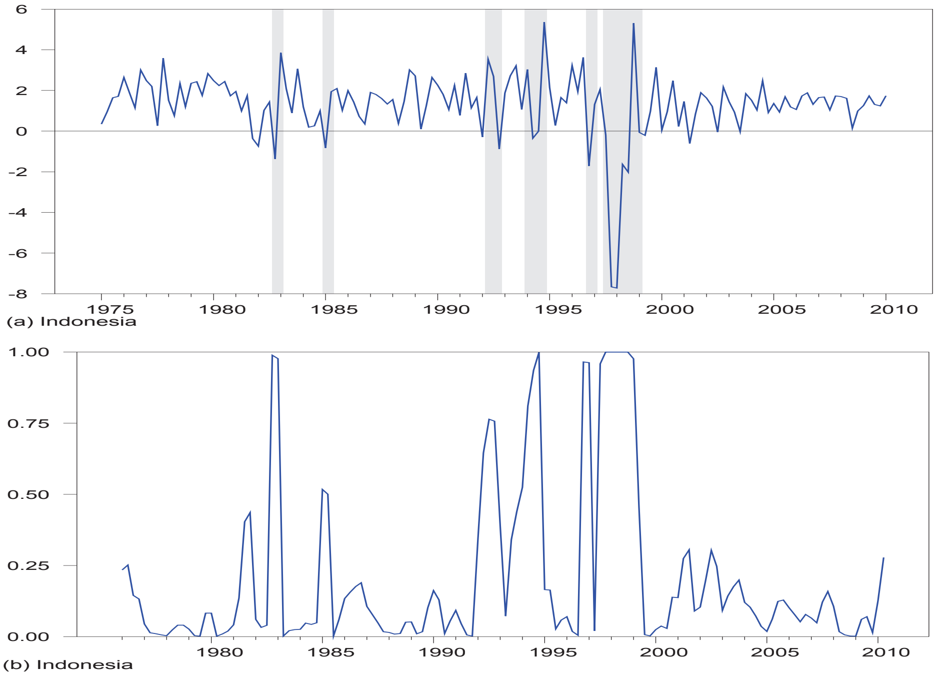
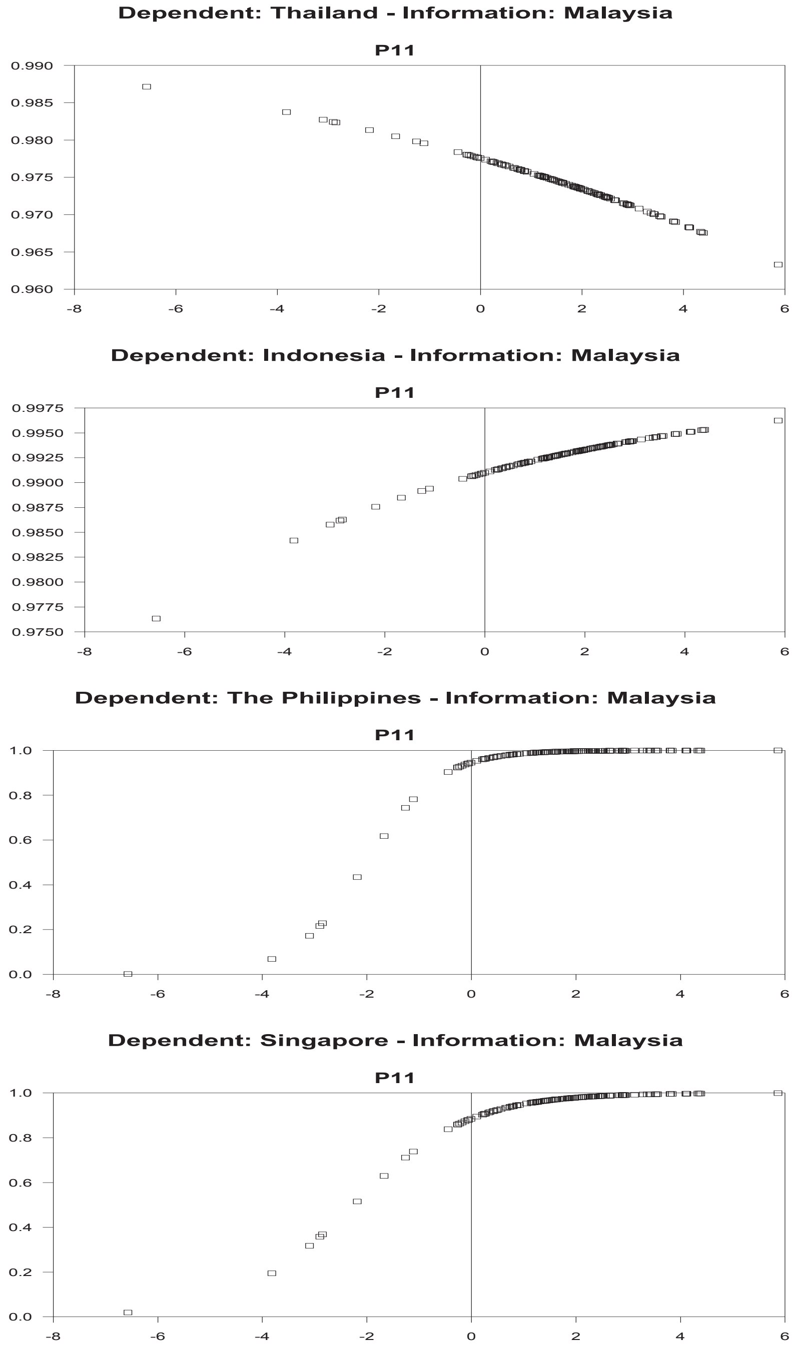


Figure 2.6: Time-varying probabilities of Southeast Asian business cycles and changes in the GDP growth of Malaysia.



Appendix 2

Table 2.7: Stylized facts on the trade integration in East Asia.

	75-80	80-85	85-90	90-95	95-00	00-05	05-10
Intra-regional trade	30,72	33,74	37,02	42,45	44,71	47,01	45,92
Share of China	9,79	11,57	17,40	19,81	23,73	27,38	30,51
Share of Japan	35,06	33,50	30,86	26,87	24,58	22,42	18,84
Total trade with US							
Singapore	14,12	14,42	19,05	17,97	17,80	14,35	10,04
Indonesia	16,72	11,83	11,33	7,51	8,30	8,43	5,87
Malaysia	15,25	14,71	16,82	17,56	18,72	18,05	13,18
Thailand	12,55	13,67	15,75	15,75	15,87	14,59	9,69
Philippines	26,25	27,56	28,59	26,54	25,90	21,83	15,53
Total trade with Japan							
Singapore	13,92	14,30	14,78	14,95	12,76	9,89	6,53
Indonesia	37,28	39,67	36,57	29,07	21,14	19,46	15,70
Malaysia	21,03	23,00	21,16	19,98	17,31	14,69	11,21
Thailand	27,37	20,33	21,71	24,19	20,19	18,93	15,84
Philippines	26,40	19,68	17,78	19,87	18,49	17,84	14,50
Total trade with China							
Singapore	1,93	2,29	3,89	2,54	3,62	6,40	10,20
Indonesia	0,00	0,65	2,21	3,48	3,94	6,05	10,33
Malaysia	2,35	1,63	2,02	2,29	2,65	6,09	10,91
Thailand	1,81	2,99	3,12	2,14	3,27	6,36	10,53
Philippines	1,72	1,93	2,40	1,32	2,14	4,02	8,78

Source: Own calculations with data from IMF DOTS.

Table 2.8: Estimations results for Malaysia as endogenous and global cycles

z	USA				Japan				China			
x	Sing.	Indo.	Thai.	Phil.	Sing.	Indo.	Thai.	Phil.	Sing.	Indo.	Thai.	Phil.
Lag k	$t-0$	$t-0$	$t-3$	$t-3$	$t-3$	$t-1$	$t-0$	$t-2$	$t-2$	$t-4$	$t-2$	$t-0$
μ_1	1,44***	1,34***	1,79***	1,77***	1,45***	1,37***	1,72***	1,77***	1,43***	1,24***	1,22***	1,78***
	(6,44)	(6,88)	(10,07)	(13,86)	(6,33)	(7,41)	(8,98)	(13,56)	(7,94)	(6,51)	(6,25)	(13,44)
μ_2	-1,63***	-2,11***	-0,63**	-3,01***	-1,4**	-1,98***	-0,75*	-3,05***	-1,42***	-2,82***	-0,71	-3,1***
	(-2,92)	(-3,80)	(-2,06)	(-8,11)	(-2,29)	(-3,92)	(-1,85)	(-5,63)	(-3,88)	(-3,66)	(-1,18)	(-9,30)
ϕ	0,07	0,15	0,01	0,06	0,05	0,12	-0,02	0,07	0,05	0,15	0,21***	0,07
	(0,67)	(1,38)	(0,05)	(0,75)	(0,62)	(1,21)	(-0,20)	(0,76)	(0,53)	(1,56)	(2,64)	(0,76)
σ	1,21***	1,14***	1,2***	1,24***	1,21***	1,16***	1,21***	1,24***	1,17***	1,17***	1,23***	1,19***
	(15,04)	(14,02)	(15,79)	(17,29)	(15,65)	(13,33)	(15,75)	(17,55)	(13,98)	(14,95)	(14,32)	(15,48)
β_1	0,2**	0,31***	0,1	0,04	0,19**	0,31***	0,12	0,04	0,18**	0,31***	0,22***	0,01
	(2,57)	(4,13)	(1,34)	(0,63)	(2,33)	(3,97)	(1,49)	(0,63)	(2,59)	(4,30)	(3,07)	(0,09)
β_2	0,99***	0,57***	0,77***	-0,04	1,07***	0,6***	0,78***	-0,05	1,12***	0,46***	1,13***	-0,06
	(5,16)	(4,59)	(6,93)	(-0,13)	(5,18)	(4,67)	(6,55)	(-0,15)	(6,60)	(2,65)	(5,46)	(-0,27)
$\theta_{1,1}$	2,28***	2,3***	3,8***	4,15***	2,43***	2,51***	3,05***	3,59***	3,76***	4,96***	2,45**	5,24***
	(3,07)	(3,63)	(3,50)	(4,69)	(3,14)	(3,67)	(3,51)	(6,03)	(2,96)	(3,42)	(2,31)	(4,35)
$\theta_{2,1}$	0,53	0,01	0,44	-0,19	-0,16	-0,34	1,14	0,48	1,64	0,46	-37,72	2,64
	(0,51)	(0,01)	(0,55)	(-0,17)	(-0,16)	(-0,49)	(1,56)	(0,66)	(0,62)	(0,31)	(0,00)	(1,21)
$\theta_{1,2}$	**1,42****	**1,12***	-1,01	-0,49	-0,09	**1,35***	**1,42****	0,43	**-0,47***	**-0,46***	-0,2	**-0,47***
	(2,01)	(1,88)	(-1,20)	(-0,72)	(-0,18)	(1,97)	(2,27)	(1,18)	(-1,75)	(-1,71)	(-0,66)	(-1,92)
$\theta_{2,2}$	-0,6	-0,76	1,22	1,15	-0,5	0,19	-0,2	0,02	-1,06	-0,39	0,04	-0,8
	(-0,67)	(-0,89)	(1,18)	(0,92)	(-0,58)	(0,59)	(-0,44)	(0,03)	(-0,75)	(-0,57)	(0,00)	(-1,10)
LR	4,81	6,16	4,05	1,49	0,54	5,54	6,11	1,10	4,55	10,28	2,77	5,09
	[0,09]	[0,05]	[0,13]	[0,48]	[0,76]	[0,06]	[0,05]	[0,58]	[0,10]	[0,01]	[0,25]	[0,08]
$Q(1)$	4,17	2,86	2,21	0,18	4,78	1,48	0,68	0,21	2,65	1,14	0,00	0,42
	[0,04]	[0,09]	[0,14]	[0,67]	[0,03]	[0,22]	[0,41]	[0,65]	[0,10]	[0,28]	[0,95]	[0,52]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.9: Estimations results for Thailand as endogenous and global cycles

z	USA				Japan				China			
x	Mal.	Sing.	Indo.	Phil.	Mal.	Sing.	Indo.	Phil.	Mal.	Sing.	Indo.	Phil.
Lag k	$t-0$	$t-0$	$t-0$	$t-0$	$t-2$	$t-4$	$t-4$	$t-3$	$t-3$	$t-1$	$t-2$	$t-1$
μ_1	1,3***	1,4***	1,43***	1,62***	1,31***	1,46***	1,52***	1,67***	1,35***	1,3***	1,55***	1,64***
	(7,39)	(8,06)	(7,21)	(12,15)	(7,15)	(8,26)	(8,10)	(11,29)	(6,70)	(7,31)	(8,11)	(10,96)
μ_2	-2,06***	-2,98***	-3,32***	-4,35***	-1,7***	-2,29***	-2,24***	-3,69***	-1,17***	-3,28***	-2,02***	-3,34***
	(-4,05)	(-5,90)	(-3,03)	(-5,11)	(-3,50)	(-5,77)	(-3,48)	(-4,81)	(-3,04)	(-5,55)	(-3,07)	(-3,82)
ϕ	-0,05	-0,03	-0,03	0,02	-0,13*	-0,14*	-0,12	0,03	-0,06	-0,07	-0,13	0,01
	(-0,51)	(-0,36)	(-0,23)	(0,21)	(-1,68)	(-1,66)	(-1,18)	(0,22)	(-0,66)	(-0,74)	(-1,36)	(0,07)
σ	1,38***	1,39***	1,43***	1,45***	1,43***	1,45***	1,38***	1,44***	1,45***	1,36***	1,37***	1,47***
	(16,35)	(17,84)	(16,37)	(16,30)	(17,14)	(16,57)	(11,55)	(15,59)	(15,63)	(14,55)	(10,84)	(14,99)
β_1	0,29***	0,2***	0,2*	0,07	0,3***	0,19**	0,22*	0,05	0,24**	0,24***	0,2	0,02
	(3,40)	(2,73)	(1,78)	(1,00)	(3,25)	(2,14)	(1,86)	(0,61)	(2,26)	(2,91)	(1,54)	(0,25)
β_2	0,69***	0,95***	0,19	0,14	0,66***	0,68***	0,36**	-0,03	0,82***	0,72***	0,39***	-0,03
	(3,91)	(4,45)	(0,94)	(0,49)	(3,92)	(5,31)	(2,58)	(-0,08)	(5,79)	(4,01)	(2,70)	(-0,09)
$\theta_{1,1}$	3,22***	3,28***	3,57***	3,77***	4,17***	3,78***	3,24***	3,68***	2,34***	3,84***	4,44***	3,67***
	(3,71)	(3,98)	(3,89)	(4,51)	(4,57)	(5,32)	(4,62)	(6,35)	(3,30)	(3,83)	(3,34)	(3,80)
$\theta_{2,1}$	-0,95	-0,91	0,04	-1,01	3,2*	1,51	0,19	-0,54	-5,2	3,66**	1,21	-0,36
	(-0,59)	(-0,61)	(0,04)	(-0,57)	(1,70)	(1,13)	(0,20)	(-0,43)	(-0,91)	(2,03)	(0,70)	(-0,31)
$\theta_{1,2}$	**1,82*** **	**2,02*** **	**1,91** **	**2,04*** **	0,05	**0,85* **	-0,02	-0,09	**0,78** **	-0,26	**-0,49* **	-0,04
	(2,63)	(2,68)	(2,51)	(2,93)	(0,06)	(1,87)	(-0,03)	(-0,18)	(2,03)	(-0,89)	(-1,91)	(-0,12)
$\theta_{2,2}$	2,26	2,18	-0,05	1,28	2,15	2,34	0,54	-1,35	5,5	**-2,17** **	-0,44	0,35
	(1,25)	(1,34)	(-0,04)	(0,69)	(1,54)	(1,21)	(0,48)	(-0,90)	(1,05)	(-2,00)	(-0,85)	(0,66)
LR	7,63	11,71	5,28	10,51	3,12	5,28	2,91	1,31	10,17	3,83	5,87	0,43
	[0,02]	[0,00]	[0,07]	[0,01]	[0,21]	[0,07]	[0,23]	[0,52]	[0,01]	[0,15]	[0,05]	[0,81]
$Q(1)$	0,75	0,00	0,05	0,36	0,23	0,31	0,08	0,01	0,03	0,03	0,01	0,00
	[0,39]	[0,99]	[0,82]	[0,55]	[0,63]	[0,58]	[0,78]	[0,92]	[0,86]	[0,85]	[0,94]	[0,98]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.10: Estimations results for Indonesia as endogenous and global cycles

z	USA				Japan				China			
x	Mal.	Sing.	Thai.	Phil.	Mal.	Sing.	Thai.	Phil.	Mal.	Sing.	Thai.	Phil.
Lag k	$t-4$	$t-2$	$t-1$	$t-4$	$t-4$	$t-1$	$t-2$	$t-1$	$t-2$	$t-2$	$t-1$	$t-1$
μ_1	1,39***	1,27***	1,29***	1,49***	1,33***	1,27***	1,33***	1,5***	1,27***	1,23***	1,29***	1,44***
	(9,77)	(9,68)	(9,37)	(16,31)	(10,20)	(13,08)	(10,34)	(15,57)	(8,89)	(9,03)	(8,87)	(14,57)
μ_2	-1,4***	-8,22***	0,3	-2,8***	-1,25***	-8,29***	-0,52	-2,77***	-1,31***	-8,29***	0,36	-2,78***
	(-4,90)	(-4,91)	(0,35)	(-7,50)	(-4,94)	(-5,37)	(-1,09)	(-7,84)	(-3,93)	(-5,07)	(0,36)	(-6,42)
ϕ	-0,08	-0,11	0,03	-0,05	-0,08	-0,12*	-0,08	-0,04	-0,08	-0,13*	0,05	-0,05
	(-1,16)	(-1,17)	(0,22)	(-0,64)	(-1,05)	(-1,96)	(-0,86)	(-0,60)	(-1,01)	(-1,84)	(0,29)	(-0,63)
σ	1,12***	1,19***	1,08***	1,11***	1,15***	1,19***	1,08***	1,11***	1,15***	1,19***	1,08***	1,12***
	(16,05)	(16,86)	(15,97)	(18,68)	(20,91)	(20,94)	(17,70)	(19,05)	(18,72)	(19,56)	(14,61)	(17,67)
β_1	0,14*	0,12**	0,09	0,02	0,13*	0,12**	0,08	0,02	0,16**	0,12**	0,06	0,02
	(1,91)	(2,01)	(1,44)	(0,32)	(1,97)	(2,54)	(1,21)	(0,25)	(2,30)	(2,01)	(0,98)	(0,36)
β_2	1,08***	-0,21	1,71***	2,93***	1,12***	-0,25	1,63***	2,77***	1,1***	-0,25	1,7***	2,96***
	(10,50)	(-0,22)	(9,96)	(8,68)	(9,64)	(-0,28)	(9,62)	(8,14)	(10,54)	(-0,26)	(9,62)	(8,24)
$\theta_{1,1}$	2,22**	5,87***	3,83***	5,52***	4,5***	4,82***	2,45***	4,59***	4,12***	4,62***	2,78***	4,31***
	(2,32)	(3,37)	(4,01)	(3,31)	(4,27)	(4,62)	(3,45)	(4,69)	(2,69)	(3,45)	(3,67)	(4,11)
$\theta_{2,1}$	-12,04	4,8	-1,27	3,97	2,26**	-40,56	-0,58	3,45	0,41	2,94	0,4	0,94
	(-0,93)	(0,92)	(-0,51)	(1,37)	(2,30)	(0,00)	(-0,57)	(1,50)	(0,23)	(1,07)	(0,16)	(0,21)
$\theta_{1,2}$	**2,57***	-0,87	-0,72	-0,62	0,09	0,19	0,72	0,7	0,06	0,19	0,15	0,25
	(1,97)	(-0,83)	(-1,18)	(-0,53)	(0,10)	(0,25)	(1,39)	(0,82)	(0,11)	(0,36)	(0,63)	(0,58)
$\theta_{2,2}$	16,83	-5,81	0,93	-1,89	-0,38	-34,73	-1,51	-3,37	1,21	-2,57	-0,47	0,48
	(1,02)	(-0,93)	(0,36)	(-0,80)	(-0,23)	(0,00)	(-1,07)	(-0,97)	(0,85)	(-1,08)	(-0,37)	(0,18)
LR	5,51	1,08	1,39	2,60	2,18	2,84	3,02	1,81	3,07	3,86	0,47	0,30
	[0,06]	[0,58]	[0,50]	[0,27]	[0,34]	[0,24]	[0,22]	[0,40]	[0,21]	[0,15]	[0,79]	[0,86]
$Q(1)$	2,95	0,41	1,27	0,01	4,05	0,41	0,97	0,05	4,22	0,56	1,17	0,06
	[0,09]	[0,52]	[0,26]	[0,94]	[0,04]	[0,52]	[0,32]	[0,83]	[0,04]	[0,46]	[0,28]	[0,80]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.11: Estimations results for Singapore as endogenous and global cycles

z	USA				Japan				China			
x	Mal.	Indo.	Thai.	Phil.	Mal.	Indo.	Thai.	Phil.	Mal.	Indo.	Thai.	Phil.
Lag k	$t-1$	$t-0$	$t-3$	$t-1$	$t-2$	$t-0$	$t-2$	$t-2$	$t-3$	$t-1$	$t-1$	$t-0$
μ_1	4,87***	1,76***	2,27***	2,16***	3,55***	1,76***	2,26***	2,16***	4,93***	1,81***	2,65***	2,22***
	(4,05)	(9,54)	(11,20)	(15,63)	(5,21)	(9,86)	(10,78)	(14,83)	(3,02)	(10,31)	(7,67)	(15,49)
μ_2	0,8***	-2,02***	-0,68*	-1,29***	0,76***	-1,98***	-0,71*	-1,27***	0,78***	-1,84***	0,51*	-1,18***
	(3,87)	(-2,73)	(-1,80)	(-2,93)	(3,43)	(-2,75)	(-1,93)	(-2,71)	(3,41)	(-2,94)	(1,94)	(-3,01)
ϕ	0,21**	0,13	-0,01	0,01	0,27***	0,14	-0,01	0,02	0,21**	0,11	0,21***	-0,02
	(2,59)	(1,03)	(-0,09)	(0,12)	(3,53)	(1,06)	(-0,01)	(0,14)	(2,54)	(0,98)	(2,80)	(-0,18)
σ	1,32***	1,39***	1,35***	1,35***	1,3***	1,4***	1,35***	1,35***	1,37***	1,44***	1,49***	1,38***
	(15,46)	(22,21)	(16,80)	(15,58)	(13,59)	(20,63)	(17,40)	(15,55)	(12,59)	(15,16)	(19,04)	(16,98)
β_1	1,14	0,24***	0,03	0,07	1,42***	0,24***	0,03	0,07	1,07	0,25***	-0,19	0,08
	(1,51)	(2,95)	(0,24)	(1,11)	(3,91)	(2,88)	(0,24)	(1,14)	(1,12)	(3,05)	(-1,60)	(1,16)
β_2	0,43***	0,52	0,33**	0,19	0,39***	0,5	0,32**	0,19	0,44***	0,66	0,59***	0,29
	(5,82)	(1,07)	(2,38)	(0,68)	(5,53)	(1,09)	(2,34)	(0,70)	(6,80)	(1,50)	(5,88)	(1,24)
$\theta_{1,1}$	3,76	2,98***	3,24***	2,88***	-2,7	3,17***	2,78***	2,95***	-4,99	2,91***	5,3***	3,47***
	(0,52)	(4,71)	(3,63)	(4,95)	(-0,95)	(5,26)	(5,56)	(5,76)	(-0,45)	(3,15)	(3,05)	(3,15)
$\theta_{2,1}$	3,83***	0,92	0,3	0,83	3,15***	0,75	1,18**	0,97	4,14***	-10,87***	3,83*	2,98*
	(4,04)	(1,32)	(0,42)	(1,46)	(3,91)	(0,86)	(2,02)	(1,51)	(2,88)	(-4,47)	(1,89)	(1,92)
$\theta_{1,2}$	-4,7	0,42	-0,45	0,32	1	0,04	0,33	0,36	1,64	0,08	**-0,61****	-0,18
	(-0,59)	(0,62)	(-0,59)	(0,54)	(0,64)	(0,09)	(0,93)	(1,07)	(0,37)	(0,22)	(-1,99)	(-0,55)
$\theta_{2,2}$	-0,26	-1,03	1,37	-0,03	**1,73****	-1,13	0,76	0,63	-0,2	**6,44*****	1,94	-0,81
	(-0,38)	(-1,35)	(1,37)	(-0,03)	(2,07)	(-1,33)	(1,43)	(1,19)	(-0,68)	(4,03)	(1,32)	(-1,54)
LR	0,38	2,42	2,91	0,28	7,53	3,83	3,15	2,76	0,65	5,04	14,99	2,47
	[0,83]	[0,30]	[0,23]	[0,87]	[0,02]	[0,15]	[0,21]	[0,25]	[0,72]	[0,08]	[0,00]	[0,29]
$Q(1)$	0,68	0,91	0,21	0,07	0,00	0,96	0,31	0,18	0,00	0,53	0,23	0,07
	[0,41]	[0,34]	[0,65]	[0,79]	[0,95]	[0,33]	[0,58]	[0,67]	[0,94]	[0,46]	[0,63]	[0,79]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) while p-values of LR and LB tests are displayed in brackets [.].

Table 2.12: Estimations results for the Philippines as endogenous and global cycles

z	USA				Japan				China			
x	Mal.	Sing.	Indo.	Thai.	Mal.	Sing.	Indo.	Thai.	Mal.	Sing.	Indo.	Thai.
Lag k	$t-1$	$t-4$	$t-4$	$t-2$	$t-3$	$t-4$	$t-3$	$t-2$	$t-1$	$t-2$	$t-2$	$t-1$
μ_1	1,04***	0,67***	0,9***	1,16***	0,83***	0,67***	0,88***	1,01***	0,75***	0,72***	0,99***	1***
	(4,81)	(4,14)	(5,81)	(8,04)	(6,05)	(3,60)	(7,01)	(6,81)	(3,73)	(5,59)	(6,64)	(5,47)
μ_2	-1,99***	-3,78**	-1,36**	-1,16**	-1,35***	-3,94***	-1,42**	-1,32**	-5**	-2,74***	-0,82*	-1,18**
	(-3,22)	(-2,49)	(-2,36)	(-2,00)	(-2,76)	(-2,70)	(-2,47)	(-2,02)	(-2,32)	(-3,50)	(-1,73)	(-2,33)
ϕ	-0,15	-0,02	-0,17**	-0,16*	-0,18**	-0,03	-0,19**	-0,17**	0,08	-0,31***	-0,28***	-0,17*
	(-1,55)	(-0,15)	(-2,10)	(-1,80)	(-2,15)	(-0,23)	(-2,60)	(-2,16)	(0,55)	(-3,80)	(-3,61)	(-1,86)
σ	1,54***	1,54***	1,67***	1,68***	1,7***	1,54***	1,7***	1,73***	1,59***	1,54***	1,58***	1,74***
	(14,42)	(19,16)	(21,44)	(15,40)	(16,49)	(19,85)	(17,58)	(17,34)	(15,30)	(14,96)	(17,67)	(18,93)
β_1	0,18**	0,24***	0,19**	0,06	0,19***	0,24***	0,18**	0,06	0,17**	0,3***	0,23**	0,06
	(2,00)	(4,10)	(2,13)	(0,81)	(2,75)	(3,02)	(2,42)	(0,81)	(1,98)	(4,72)	(2,56)	(0,58)
β_2	0,44*	-0,14	-0,05	0,13	-0,05	-0,09	0,07	-0,05	0,33	0,67***	-0,01	-0,01
	(1,75)	(-0,27)	(-0,15)	(0,44)	(-0,15)	(-0,18)	(0,19)	(-0,11)	(0,27)	(2,70)	(-0,01)	(-0,02)
$\theta_{1,1}$	2,03***	2,97***	12,65**	2,06*	5,02***	3,68***	5,12***	4,93***	3,06***	1,76*	1,07	3,64***
	(3,02)	(4,51)	(2,03)	(1,91)	(3,29)	(4,80)	(3,41)	(3,77)	(5,37)	(1,93)	(0,98)	(2,75)
$\theta_{2,1}$	-13,52	-75,62	0,67	-0,93	4,13*	-35,42	4,31**	3,28	-28,3	0,55	0,4	0,81
	(-1,09)	(0,00)	(0,81)	(-0,69)	(1,83)	(0,00)	(2,09)	(1,11)	(0,00)	(0,35)	(0,46)	(0,48)
$\theta_{1,2}$	0,71	0,36	**7,55****	4,31	-0,45	-0,51	-0,43	-0,32	0,11	**2,12****	**2,39***	0,57
	(1,21)	(0,64)	(2,02)	(1,47)	(-0,41)	(-1,16)	(-0,37)	(-0,35)	(0,52)	(2,33)	(1,95)	(1,34)
$\theta_{2,2}$	13,75	0,05	1,29	2,09	-1,7	0,06	-1,73	-1,01	0,02	-0,07	0,35	0,35
	(1,04)	(0,00)	(1,42)	(1,56)	(-1,33)	(0,00)	(-1,45)	(-0,60)	(0,00)	(-0,18)	(0,86)	(0,63)
LR	1,76	3,60	13,85	3,30	3,65	4,36	1,98	3,41	0,17	8,57	7,17	1,75
	[0,42]	[0,17]	[0,00]	[0,19]	[0,16]	[0,11]	[0,37]	[0,18]	[0,92]	[0,01]	[0,03]	[0,42]
$Q(1)$	1,56	1,26	0,49	0,11	1,35	1,12	0,74	0,42	3,00	0,17	0,03	0,94
	[0,21]	[0,26]	[0,48]	[0,75]	[0,24]	[0,29]	[0,39]	[0,52]	[0,08]	[0,68]	[0,86]	[0,33]

Notes: *,**,*** denote significance at 10, 5 and 1 % respectively. Student-t statistics of parameters are reported in parentheses (.) and p-values of LR and LB tests are displayed in brackets [.].

Table 2.13: Summary of the relationships among the East Asian countries and the global cycles

z	USA					Japan					China				
x	Mal.	Thai.	Phil.	Sing.	Indo.	Mal.	Thai.	Phil.	Sing.	Indo.	Mal.	Thai.	Phil.	Sing.	Indo.
Mal.	-	No	No	Yes	Yes	-	Yes	No	No	Yes	-	No	Yes	Yes	Yes
Thai.	Yes	-	Yes	Yes	Yes	No	-	No	Yes	No	Yes	-	No	Yes	Yes
Phil.	No	No	-	No	Yes	No	No	-	No	No	No	No	-	Yes	Yes
Sing.	No	No	No	-	No	Yes	No	No	-	No	No	Yes	No	-	Yes
Indo.	Yes	No	No	No	-	No	No	No	No	-	No	No	No	No	No

Notes: These results are based on the significance of coefficients $\theta_{1,2}$ and $\theta_{2,2}$.

Chapter

3

Southeast Asian Monetary Integration: New Evidences from Fractional Cointegration of Real Exchange Rates

3.1 Introduction

This empirical paper examines the long-run relationship among the RERs of the five most advanced members of the ASEAN countries.[1] Our aim is to investigate whether the ASEAN-5 has operated as an Optimum Currency Area (OCA) over the period 1975-2011. Prior research on cointegration among RERs has focused on the traditional $I(1)/I(0)$ cointegrating relationships. In the former case, there is no equilibrium relationship and related studies conclude against the existence of an OCA (Wilson and Choy , 2007; Sun and Simons , 2011), while in the latter case, all RERs share the same kind of real disturbances, thus strengthening the case of a currency union (Ogawa and Kawasaki , 2008; Mishra and Sharma , 2010). However, classical cointegration models used in these studies imply short memory residuals as the equilibrium errors are restricted to be I(0). In this paper we argue that such conflicting evidences may be resolved in a fractionally cointegrated framework.

Considering the above limitation, we depart from the traditional framework by supposing the presence of long memory in the cointegrating residuals. The rationale for considering RERs as a fractional cointegration process has its counterparts in the literature. The recent floating exchange rate period has been characterized in many countries by significant and persistent misalignment of RER. This perception has spurred a renewed interest in the study of exchange rate behavior while, at the same time, long memory models as well as their applications have become popular among researchers. Initially, Booth et al. (1982) apply a classical rescaled range analysis to examine three major daily spot exchange rates, namely the British pound, French franc and Deutsche mark. These authors find evidences of positive long-term persistence during the flexible exchange rate period (1973-1979) but negative long-term persistence during the fixed exchange rate period (1965-1971). Later, Cheung (1993), also find convincing evidences of long-term persistence in five major weekly nominal exchange rates during the managed

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1. The ASEAN-5 consists of founder members namely Singapore, Thailand, the Philippines, Indonesia, Malaysia.

floating regime, using the Geweke-Porter-Hudak test. Baum et al. (1999) estimate an ARFIMA model to test for the existence of long memory in international inflation rates (CPI- and WPI-based inflation rates) for a number of industrial and developing countries, over the post-Bretton Woods (1971-1995) period. The authors demonstrate that long memory is a common feature for the inflation rates of the countries studied.

More recent papers have shown that RERs are subject to persistent shocks, taking them away from their fundamental value over unexpected timescales. In other words, mean reversion occurs but deviations from equilibrium are long-lived (see, among others, Gil-Alana , 2000; Caporale and Gil-Alana , 2004 and Dufrénot et al. , 2008). The presence of long memory in currency movements led researchers to investigate fractional cointegration techniques for modeling long-run relationship among exchange rates. For instance, Baillie and Bollerslev (1994) argue that long-run relationship among a set of seven exchange rates may well be tied together through a long memory $I(d)$-type process. Cheung and Lai (1993) analyze Purchasing Power Parity (PPP) and find evidence that relative prices are characterized by a fractionally cointegrated process. [2]

In view of the preceding remarks, this paper seeks to test G-PPP hypothesis introduced by Enders and Hurns (1994) by means of a robust estimation methodology of fractional cointegration. We envisage in this study both US dollar-based and yen-based RERs for the ASEAN-5 countries using monthly data over the period 1975-2011 ($n = 444$). [3]

The main results of the paper are the following: 1) We find different sub-groups of the ASEAN-5 countries sharing with each other a long-run relationship. 2) All of the cointegration relationships are weak in the sense that RERs deviate persistently from the long-run equilibrium. As expected, these two results imply important limitations for a monetary union involving all the ASEAN-5 countries. 3) All significant cointegration relationships exhibit a positive long-run coefficient, thus implying symmetric adjustments to macroeconomic disturbances. 4) Both the yen and the US dollar are important when considering a step-by-step approach toward the coordination of monetary policy. 5) There are some evidences to assert that the ASEAN-5 countries became more integrated after 1997-98 since the equilibrium errors respond more rapidly to shocks, indicating less persistent deviations toward the common equilibrium in the post-crisis period.

The remainder of the paper is laid out as follows. Section 3.2 provides an overview of previous studies and clarifies the position of our analysis. Section 3.3 presents the data and introduces our empirical methodology. In Section 3.4, we discuss the results and their economic implications. Section 3.5 concludes the paper.

3.2 Monetary integration in Southeast Asia and G-PPP

After the 1997-98 crisis, Asian authorities stated their intention to create a Common Currency Area (CCA). [4] Arguments in favor of monetary integration are commonly related to the preponderance of

2. PPP theory is a structural model of RER determination supposing that two currencies in equilibrium lead to the same purchasing power in both countries.

3. Indeed, the United States is one of the most important export markets for final outputs produced inside the region through vertical integration of production chains, while the Japanese economy accounts for a large proportion of trade among Southeast Asian countries.

4. At the 39th Annual Meeting of the ADB, a first step was initiated in this direction with the introduction of the Asian Currency Unit (ACU) as a benchmark to monitor movements in the value of currencies in the region. More recently, the prospect of launching a single currency was put forward by the Japanese prime minister on 23 October 2009, during the

market-driven forces, such as trade and FDI, fostered by the constitution of an intense intra-industry trade networking (Rana , 2007). Some studies have suggested that the US dollar will be better suited for East Asian countries to guarantee implicit coordination of exchange rates movements and then anchor domestic price levels within the region (McKinnon and Schnabl , 2004). The creation of a yen block has also been envisaged by Kwan (2001) following the announcement of Japanese authorities to promote the yen's internationalization, while the possibility of a CCA based on the Chinese yuan is still limited (Park , 2010).

Monetary unions are studied through the lens of the OCA theory. The seminal paper of Mundell (1961) argued that the mobility of the labor force within an union and the flexibility of prices/wages could be considered as alternatives to flexible exchange rates.[5] However, if countries face symmetric shocks, then the need for alternative mechanisms is considerably reduced. Related researches as Bayoumi and Eichengreen (1994), Chow and Kim (2003), Obiyathulla (2008) and Genberg and Siklos (2010), using structural vector autoregression (SVAR), have shown mixed results. This literature does not necessarily tell the whole story about the relevance of an OCA within a region.[6] Indeed, the SVAR model includes first-differenced variables, thus removing any information about the long-run equilibrium relationship for a set of real variables. This can pose a greater obstacle for monetary integration if macroeconomic fundamentals among the economies are not cointegrated. A convenient complementary approach is the concept of G-PPP, as proposed by Enders and Hurns (1994). The G-PPP model was originally suggested to overcome the empirical failure of PPP in adequately explaining RER behavior.[7] According to the authors, the failing of PPP is mainly due to the continuous effect of real and nominal shocks on variables that drive RERs and, as a consequence, imply permanent movement in RERs.[8]

Subsequently, Enders and Hurns (1997) suggest that "despite the fact that domestic economies differ, the trends in real forcing variables affecting real exchange rates might be similar across diverse countries" in such a way that RERs may share long-run equilibrium relationships with each other. Indeed, PPP deviations in the long run might imply that non-stationary RERs move together consistently with common trends in both domestic and international environments. This means that a shock in any one RER within a group of two or more countries will affect other RERs, thus supporting the case of an OCA in the sense of Mundell (1961).

It follows that several recent papers have proposed examining whether the East Asian currencies shared a common trend. Using RERs of the ASEAN-5 countries, the empirical results of Wilson and Choy (2007) do not support the evidence of a CCA with either the US or Japan, while Mishra and

15th summit of the Association of Southeast Asian Nations. The economic regionalism also led to substantial reforms such a regional economic surveillance process framed in the Economic Review and Policy Dialog process (ERPD, May 2000) and a liquidity support arrangement under the form of bilateral swaps (Chiang Mai Initiative, May 2000).

5. As a monetary policy cannot be tailored to absorb country-specific shocks within a monetary union, two countries may find an advantage in adopting the same currency if there are alternative mechanisms of adjustment that can help them to cope with asymmetric shocks. See de Grauwe (2012) for a survey of the OCA theory.

6. For different empirical methods, see Girardin (2004) who uses a Markov-switching VAR, and Allegret and Essaadi (2011) who introduces a spectral analysis based on time-varying coherence function.

7. From a empirical perspective, PPP holds if bilateral RER is stationary involving RER convergence to its equilibrium in the long run. Numerous empirical studies have shown that RERs are non-stationary in level, thereby calling into question the validity of PPP as a structural model of RER determination (Alder and Lehman , 1983, Enders , 1988). However, as mentioned above, other studies, based on fractional differencing analysis, provided evidence of mean reversion in RER supporting PPP as a long-run phenomenon. In the context of Asian countries, see, among others, Bec and Zeng (2013), Mastuki and Sugimoto (2013) and, in the context of industrialized countries, Gil-Alana (2000).

8. If macroeconomic fundamentals that determine RERs, such as real output, expenditure patterns and money supply, are, themselves, non-stationary, then RERs are also non-stationary. This finding is consistent with many structural models of exchange rate determination including Dornbush (1976) and Frankel (1983).

Sharma (2010) find the opposite for seven ASEAN countries and India. Rangkakulnuwat et al. (2010) extend G-PPP theory with foreign variables such as the real money supply, the output and interest rate. They further found that selected Asian countries form a CCA with Japan as the base country. Choudhry (2005) conducted Johansen's cointegration test and found a substantial change in relationships between several RERs of the Far East Asian countries with ample evidence of G-PPP after the crisis period. Sun and Simons (2011) find a "pentagon" group of five countries (South Korea, the Philippines, Thailand, Malaysia and Indonesia) with potential success for further monetary integration even if neither a yen bloc nor a US dollar bloc is forming in East Asia. Ogawa and Kawasaki (2008) investigate the possibilities of an OCA in East Asia and find that the ASEAN+3 countries are forming an OCA, notably in the post-crisis period. The authors suggest a common exchange rate policy with reference to a common currency basket composed of the US dollar, the euro and the yen.

The above line of empirical research neglects the fact that RERs probably converge slowly toward a long-run equilibrium, implying that the autocovariance function of the cointegrating residuals declines at a slower hyperbolic rate rather than the classical geometric rate found in weak dependent cointegrating residuals (see the next section). To overcome this conceptual drawback, we propose a new empirical methodology based on fractional cointegration.

3.3 Data and empirical methodology

3.3.1 Description of the data

Monthly data include the consumer price index (CPI) and the nominal bilateral exchange rates against two numeraires, the US dollar and the Japanese yen, for five Southeast Asian countries: Malaysia (Mala.), Indonesia (Indo.), Singapore (Sing.), the Philippines (Phil.) and Thailand (Thai.). The bilateral RERs against the US dollar (USD) and the yen (YEN) are calculated according to the following formulas:

$$q_t = s_t + p_t^* - p_t$$

where q_t is the bilateral RER, s_t is the nominal exchange rate (expressed as the domestic currency price of the base currency), p_t^*, the base country price index and p_t, the domestic price index. All data were collected from International Monetary Fund *International Financial Statistics* database between 1975:1 and 2011:12. For estimation, we take the natural logarithm of all variables.

Since the crisis may have permanently affected macroeconomic fundamentals of Southeast Asian economies, we apply the Bai-Perron methodology (Bai and Perron , 2003) for identifying dates of structural change. The structural change analysis is performed on US dollar-based RERs and the results are presented in Table 3.1.[9] The break dates are estimated at 1997:11, 1997:08, 1998:05, 1997:07 and 1997:08 for the Philippines, Indonesia, Singapore, Thailand and Malaysia, respectively. This finding confirms the presumption that the main shock experienced by the Southeast Asian countries within the sample period is the Asian financial crisis.[10] Accordingly, the sample is divided into two sub-samples, the pre-crisis period (1975:1-1997:5) and the post-crisis period (1998:6-2011:12). We exclude the crisis-period (1997:6-1998-5) to the avoid possible influence of the financial crisis. The comparative analysis

9. We test the hypothesis of no break versus one break on the individual series.
10. Similar results are provided in Mastuki and Sugimoto (2013).

may be useful to assess progress toward achieving greater rate of convergence over the last decade. The investigation is also performed on the full sample.

Table 3.1: Test results of the structural changes test relative to Southeast Asian RERs

	The Philippines		Indonesia		Singapore		Thailand		Malaysia	
Breaks	BIC	LWZ	BIC	LWZ	BIC	LWZ	BIC	LWZ	BIC	LWZ
0	-3.01	-3.03	0.23	0.25	-4	-3.98	-2.65	-2.63	-1.71	-1.69
1	-3.45	-3.48	-0.55	-0.52	-4.19	-4.15	-3.7	-3.67	-3.27	-3.24
Sup$F_T(1\|0)$	267.54		541.4		99.93		846.81		1691.91	
$\hat{\eta}_1$	1.128		1.683		1.257		1.147		1.421	
	(0.01)		(0.045)		(0.007)		(0.009)		(0.012)	
$\hat{\eta}_2$	1.4045		3.381		1.377		1.586		2.193	
	(0.013)		(0.057)		(0.009)		(0.012)		(0.015)	
$\hat{T}_1$	97:11		97:08		98:05		97:07		97:08	
	(94:05-99:07)		(95:11-98:04)		(91:12-04:08)		(97:05-98:03)		(97:05-98:03)	

Notes: The table presents structural break estimations on individual series, using the method of Bai and Perron (1998, 2003). We consider the case of a pure structural change model with $m = 1$ break. The regression is given by: $y_t = x_t'\eta_j + u_t; t = T_{j-1} + 1, \cdots, T_j$ for $j = 1, \cdots, m+1$. In this model, y_t is the observed dependent variable at time t; x_t is a vector $(p \times 1)$ of covariates and $\eta_j (j = 1, \cdots, m+1)$ is the corresponding vector of coefficients; u_t is the disturbance at time t. The indices $(T_1, \cdots, T_m)$ are the break points. We apply the procedure with only a constant as regressor (i.e. $x_t = \{1\}$). The null hypothesis of $SupF_T(1)$ test is no structural break ($m = 0$) and the alternative hypothesis is $m = 1$ structural break. The Sup$F_T(1|0)$ BIC and the modified Schwarz criterion of Liu et al. (1997) select one break. In parentheses are the standard errors for $\hat{\eta}_i (i = 1, 2)$ and the 95% confidence interval for the estimated break point.

3.3.2 Strategy of estimation and testing

In this paper, we conduct our analysis using a bivariate fractional cointegration approach that is a generalization of the traditional $I(1)/I(0)$ cointegration in the sense that integration orders of the regressor (δ) and the cointegrating residuals (γ) are real numbers and not just integers.[11] Considering this framework, two RERs series, $q_{t,a}$ and $q_{t,b}$, are said to be cointegrated if $q_{t,a}$ and $q_{t,b}$ are $I(\delta)$ and for some real $\xi > 0$, there exits a coefficient β such that the linear combination of $q_{t,a}$ and $q_{t,b}$ is $I(\gamma = \delta - \xi)$. This general definition includes cases that traditional cointegration estimators cannot handle because they assume weak dependent cointegrating errors.

Fractional cointegration can generally be investigated following two different methods (see Gil-Alana and Hualde , 2009 for an overview). The first is the so-called regression-based approach and requires to estimate the cointegrating vector(s), the integration orders of the regressor(s) and the cointegrating residuals. The second consists of estimating the cointegrating rank and does not imply any such requirement. In both methodologies, one can encounter some difficulties to obtain entirely conclusive results. For instance, it is sometimes tedious to test for cointegration in the regression-based approach when integration orders are not confined in a particular region of the parameter space or when the resid-

11. A multivariate approach is possible using the methodology of Johansen and Nielsen (2012). Nonetheless, as emphasized by Sun and Simons (2011), joining a monetary union is a country-level decision and therefore, a bivariate setting is more appropriated.

uals are not well defined as in semi-parametric methods. In such cases, there is little consensus on how to test for fractional cointegration. On the other hand, rank analysis often requires tuning parameters that increases the sensitivity of the results (see Robinson and Yajima , 2002) and does not bring information on the strength of the relation. To overcome these difficulties and to be more confident in the interpretation of the results, we adopt the two strategies. Therefore, our methodology relies on three steps which consists of testing for the homogeneity of the integration orders of pairwise RERs, estimating the cointegrating rank and finally estimating the long-run relationship. [12]

3.3.3 Homogeneity of integration orders and G-PPP

A first requirement for two RERs to be non-trivially cointegrated is the homogeneity of their integration orders. Therefore, in a first step, we test for the null hypothesis of equality of the integration orders, $H_0 : \delta_a = \delta_*, a = 1, \ldots, p, p = 2$, by applying a test originally proposed by Robinson and Yajima (2002) and extended to non-stationary processes by Nielsen and Shimotsu (2007). This test requires preliminary estimates of long memory parameters. We employ here the two-step exact local Whittle (2S-ELW) estimator of Shimotsu (2010) because it accommodates both, stationarity and non-stationarity and is invariant to polynomial time trend and unknown mean. More precisely, the author proposes to deal with the polynomial trend using the tapered estimator of Velasco (1999) in a first stage and to estimate the unknown mean, μ_a, of the series $q_{t,a}$, by $\hat{\mu}_a = w(\delta_a)\bar{q}_a + (1 - w(\delta_a))q_{1,a}$, where $\bar{q}_a = n^{-1} \sum_{t=1}^{n} q_{t,a}$ and $w(\delta)$ is a smooth weight function such that $w(\delta) = 1$ for $\delta \leq 1/2$, $w(\delta) \in [0, 1]$ for $1/2 \leq \delta \leq 3/4$ and $w(\delta) = 0$ for $\delta \geq 3/4$. In a second stage, the author estimates δ_a by $\hat{\delta}_a = \arg \min_{\delta \in [\Delta_1, \Delta_2]} R(\delta_a)$ where

$$R(\delta_a) = \log G_a(\delta_a) - 2\delta_a \frac{1}{m} \sum_{j=1}^{m} \log \lambda_j, \qquad G_a(\delta_a) = \frac{1}{m} \sum_{j=1}^{m} I_{\Delta^\delta(q_a - \hat{\mu}_a)}(\lambda_j), \tag{3.1}$$

with Δ_1 and Δ_2 the lower and upper bounds of the admissible values of δ_a, $\lambda_j = (2\pi j/n)$ the fundamental frequency and $I_{\Delta^\delta(q_a - \hat{\mu}_a)}$, the periodogram of $\Delta^\delta(q_a - \hat{\mu}_a)$. Following the recommendation of Robinson (2008b), we investigate a grid of values for the bandwidth $m = n^k$ ($k = \{0.5, 0.6, 0.7, 0.8\}$).

Table 3.2: 2S-ELW preliminary estimates of δ - 1975:01 to 2011:12 ($n = 444$)

USD					
Bandwidths	Indo.	Mal.	Phil.	Sing.	Thai.
0,5	0.7850	0.9106	1.0131	1.1761	0.9210
0,6	0.9362	1.0205	1.1123	1.1050	0.9761
0,7	0.9699	1.0857	1.0779	1.0821	1.0168
0,8	1.0412	1.1510	1.0807	1.1401	1.1285
YEN					
Bandwidths	Indo.	Mal.	Phil.	Sing.	Thai.
0,5	0.8634	0.9799	0.8498	1.0691	0.9667
0,6	1.0459	1.0414	0.9850	1.1577	0.9089
0,7	0.9927	1.0919	1.0426	1.1523	1.0236
0,8	1.0110	1.1884	1.0857	1.1816	1.1341

Results in Table 3.2 are not so sensitive to bandwidth selection. Unfortunately, this is not the case in the rest of the paper (i.e. concerning the rank analysis and the regression-based approach). In such

12. A similar methodology can be found in Nielsen and Frederiksen (2011).

situation, we support that a small bandwidth is preferable. Indeed, when the bandwidth parameter is large, estimates are more likely to be contaminated by high frequency, in presence of short-run dynamics. Conversely, when the bandwidth parameter is small, estimates are more likely to face large variance. Given this trade-off and the sample size, we choose to only report the results for $m = n^{0.6}$ in the next tables. [13]

The test statistic proposed by Robinson and Yajima (2002) is given by

$$\hat{T}_0 = m(S\hat{\delta}_{ab})' \left(S\frac{1}{4}\hat{D}^{-1}(\hat{G} \odot \hat{G})\hat{D}^{-1}S' + h(n)^2 I_{p-1} \right)^{-1} (S\hat{\delta}_{ab}), \qquad p = 2, \tag{3.2}$$

where $\hat{\delta}_{ab} = (\hat{\delta}_a, \hat{\delta}_b)'$, $h(n) > 0$, I_p denotes the identity matrix of dimension p, $\odot$ denotes the Hadamard product, $\hat{D} = \mathrm{diag}\left\{G_{11}, \dots, G_{pp}\right\}$ and $S = [I_{p-1}; -\iota]$, with ι a $(p-1)$-vector of ones. Nielsen and Shimotsu (2007) argue that the test statistic in Eq. (3.2) holds with non-stationary data. Under null hypothesis, $\hat{T}_0 \xrightarrow{d} \chi^2_{p-1}$ if $q_{t,ab}$ is not cointegrated and $\hat{T}_0 \xrightarrow{p} 0$ if $q_{t,ab}$ is cointegrated, with $q_{t,ab} = (q_{t,a}, q_{t,b})'$. The matrix $\hat{G}$ is obtained from a multivariate extension of the estimator of Shimotsu (2010). [14] The test statistic is calculated with $h(n) = (\log n)^{-1}$.

Table 3.3: Test of homogeneity for USD and YEN - $m = n^{0.6}$ - 1975:01 to 2011:12 ($n = 444$)

	Indo.	Mal.	Phil.	Sing.	Thai.
Indo.		0.566	2.404	1.827	0.130
		(0.452)	(0.121)	(0.176)	(0.719)
Mal.	0.002		0.690	0.683	0.196
	(0.968)		(0.406)	(0.408)	(0.658)
Phil.	0.311	0.418		0.004	1.446
	(0.577)	(0.518)		(0.953)	(0.229)
Sing.	0.806	1.795	3.394		1.277
	(0.369)	(0.180)	(0.065)		(0.258)
Thai.	1.766	2.243	0.663	5.770	
	(0.184)	(0.134)	(0.416)	(0.016)	

Notes: P-values are displayed in parentheses (.).

Preliminary estimates of δ_a reveal that most of integration orders are confined in a close interval around unity (see Table 3.2). Therefore, it is obvious that RERs are not weak dependent (i.e. short memory) and according to Enders and Hurns (1994), if the RERs possesses a unit root, we can apply the G-PPP theory. Table 3.3 reports the test statistic $\hat{T}_0$ for the 2S-ELW estimates of the pairwise fractional integration orders. If the null hypothesis is accepted at conventional significance levels, the pairwise integration orders are equals. Since the 95% critical value of the χ^2_1 distribution is 3,84 we easily accept, in most cases, the null of equality of the integration order. We also investigate whether the East Asian RERs are defined as unit root ($\delta_a = 1$) or persistent mean reverting processes ($\delta_a < 1$) using the fractional unit root test of Lobato and Velasco (2007), henceforth LV. The LV fractional Wald test consists of restating $(1-L)^{\delta_a} q_{t,a} = \varepsilon_t$, to test whether δ_a is not significantly different from 1 under the null hypothesis. In other words, we test $q_{t,a}$ as a unit root process against the fractional alternative:

13. All results are computed using Matlab 7.9 or RATS 8.1. Unreported results are available upon request.

14. Notice that in presence of cointegration, the matrix G has reduced rank. Nonetheless, the 2S-ELW makes no assumptions about the presence or absence of cointegration and Nielsen and Shimotsu (2007) show its consistency in both cases.

$$(1-L)q_{t,a} = \varphi\zeta_{t-1}(\delta_a) + \varepsilon_t, \qquad \zeta_{t-1}(\delta_a) = \frac{(1-L)^{\delta_a - 1} - 1}{1-\delta_a}(1-L)q_{t,a} \tag{3.3}$$

The LV test also requires to pre-estimate of δ_a with a consistent estimator. We employ the results obtained from Shimotsu (2010) for $m = n^{0.6}$. Table 3.4 reports the full sample results of the fractional unit root test and highlights that all examined exchange rates possess a unit root. Therefore, the results confirm that the G-PPP theory is appropriated.

Table 3.4: LV test based on the 2S-ELW ($m = n^{0.6}$) - 1975:01 to 2011:12 ($n = 444$)

	q_t	2S-ELW	Stat.	Signif.
USD	Indo.	0.9362	0.3679	0.7128
	Mala.	1.0205	-0.9710	0.3315
	Phil.	1.1123	0.6260	0.5313
	Sing.	1.1050	-0.8564	0.3917
	Thai.	0.9761	0.0183	0.9853
YEN	Indo.	1.0459	0.7180	0.4727
	Mala.	1.0414	0.3907	0.6960
	Phil.	0.9850	-0.6318	0.5274
	Sing.	1.1577	0.8170	0.4139
	Thai.	0.9089	-0.2800	0.7794

3.3.4 The rank analysis

The second step of our methodology consists of performing the rank analysis of pairwise RERs using the rank estimator of Nielsen and Shimotsu (2007). This procedure requires to estimate the common value of δ_a and δ_b, denoted δ_*. Although the previous tests conclude to homogeneity of parameters, they do not provide information on δ_*. An intuitive approach to estimate δ_* would be to apply a constraint estimation of the bivariate version of the 2S-ELW for each pairs of series. Unfortunately, this method is unfeasible in presence of cointegration because $G(\delta_*)$, defined as $G(\delta_*) = \frac{1}{m_1}\sum_{j=1}^{m_1} I_{\Delta(L;\delta_*,\delta_*)r}(\lambda_j)$ where $\Delta(L;\delta_a,\delta_b)r$ denotes the vector $(\Delta^{\delta_a}q_{t,a}, \Delta^{\delta_b}q_{t,b})'$, does not have full rank in this case. Following Nielsen and Shimotsu (2007), we trivially estimate the common value δ_* by $\bar{\delta}_* = p^{-1}\sum_{a=1}^{p}\hat{\delta}_a, p = 2$. Then we estimate $G(\delta_*)$ by substituting δ_* with $\bar{\delta}_*$. As mentioned in Nielsen and Shimotsu (2007), $\bar{\delta}_*$ also needs to converge to δ_* at a faster rate than $m_1^{1/2}$. Therefore, we accommodate this convergence rate requirement using m ordinates in estimating δ_a and δ_b, with $m/m_1 \to 0$ and m_1 ordinates in estimating $G(\bar{\delta}_*)$, with $m_1 = m - 0.05$. Empirical mean of pairwise estimates are reported in Table 3.5.

Table 3.5: Empirical estimates of δ_* for USD and YEN - 1975:01 to 2011:12 ($n = 444$)

δ_*	Indo.	Mal.	Phil.	Sing.	Thai.
Indo.		0.978	1.024	1.021	0.956
Mal.	1.044		1.066	1.063	0.998
Phil.	1.015	1.013		1.109	1.044
Sing.	1.102	1.100	1.071		1.041
Thai.	0.977	0.975	0.947	1.033	

We conduct the rank analysis considering both, the matrix $\hat{G}(\bar{\delta_*})$ and the correlation matrix $\hat{P}(\bar{\delta_*}) = \hat{D}(\bar{\delta_*})^{-1/2}\hat{G}(\bar{\delta_*})\hat{D}(\bar{\delta_*})^{-1/2}$ with $\hat{D}(\bar{\delta_*}) = \mathrm{diag}\left\{\hat{G}_{11}(\bar{\delta_*}), \ldots, \hat{G}_{pp}(\bar{\delta_*})\right\}, p = 2$. However, Nielsen and Shimotsu (2007) document a Monte Carlo experiment that reveals that selection procedure based on $\hat{P}(\bar{\delta_*})$ performs better. Therefore, we only report the results with $\hat{P}(\bar{\delta_*})$. Denoting τ_i the i-th eigenvalue of $\hat{G}(\bar{\delta_*})$ or $\hat{P}(\bar{\delta_*})$, the rank estimate is defined as

$$\hat{r} = \arg \min_{u=0,\ldots,p-1} L(u), \qquad L(u) = \upsilon(n)(p-u) - \sum_{a=1}^{p-u} \tau_a, \qquad p = 2 \tag{3.4}$$

Equation of $L(u)$ embed an additional tuning parameters, $\upsilon(n)$, for which the procedure is highly sensitive. In our application, we employ $\upsilon(n) = m_1^{-k}$ with $k = \{0.35, 0.25, 0.15\}$. Simulations of Nielsen and Shimotsu (2007) reveal that for high values of k, the procedure is more likely to provide underestimates of $\hat{r}$ while for small values of k, the procedure is more likely to provide overestimate of $\hat{r}$.

3.3.5 The regression-based approach

The rank analysis is convenient to test for cointegration but reveals only few information on the strength of the long-run relationship. Conversely, the regression-based approach provides such useful information although it involves some complication in testing for cointegration. We perform here the regression-based approach to confirm the rank estimation and to go further in the cointegration analysis. We consider the following bivariate model,

$$(1-L)^{\gamma}(q_{t,a} - \beta q_{t,b}) = \varepsilon_{1t}, \qquad (1-L)^{\delta} q_{t,b} = \varepsilon_{2t}, \qquad t = 1, 2, ..., n, \tag{3.5}$$

where $q_{t,a}$ and $q_{t,b}$ are two different exchange rates and $\delta = \delta_a = \delta_b$. The fractional filter, $(1-L)^{\alpha}$, is defined by the following binomial expansion

$$(1-L)^{\alpha} = \sum_{k=0}^{+\infty} a_k(\alpha) L^k, \qquad a_k(\alpha) = \frac{\Gamma(k-\alpha)}{\Gamma(k+1)\Gamma(-\alpha)} \tag{3.6}$$

where $\Gamma(z) = \int_0^{+\infty} t^{z-1} e^{-t} dt$ is the Gamma function.

From the model in Eq. 3.5, the literature generally identifies three subspaces, closely related to the strength of the cointegration relationship: the strong fractional cointegration ($\delta - \gamma > 1/2$); the weak fractional cointegration ($\delta - \gamma < 1/2$, $\delta > 1/2$); the stationary fractional cointegration ($\delta - \gamma < 1/2$, $\delta < 1/2$). According to our preliminary results, we are likely to found weak or strong fractional cointegration relationships. One difficulty concerning this approach is that most of estimators are not able to provide consistent estimates of β, δ and γ for the three aforementioned subspaces. With respect to our bivariate framework, the semi-parametric exact local Whittle estimator of fractional cointegration (2S-ELW-FC) of Shimotsu (2012) is particularly appropriated. This estimator jointly estimates all the parameters of interest, β, γ and δ (henceforth denoted θ), and operates in two-steps in order to accommodate both stationary and non-stationary processes for the stochastic trend and the cointegrating error. The first step consists of an estimation of the tapered local Whittle (LW) estimator of Robinson (2008a). Denoting $z_t = (q_{t,a} - \beta q_{t,b}, q_{t,b})'$ and $\theta^1 = (\gamma, \delta)'$, the objective function of the tapered LW estimator is defined by,

$$R(\theta) = \log \det \hat{\Omega}(\theta) - 2(\gamma + \delta) \frac{g}{(1-\kappa)m} \sum_{j(g,\kappa)}^{m} \log \lambda_j, \tag{3.7}$$

$$\hat{\Omega}(\theta)\frac{g}{(1-\kappa)m} = \sum_{j(g,\kappa)}^{m} \mathrm{Re}\left[\Psi(\lambda_j;\theta^1)BI_z(\lambda_j)B'\Psi(\lambda_j;\theta^1)^*\right], \qquad B = \begin{pmatrix} 1 & -\beta \\ 0 & 1 \end{pmatrix} \tag{3.8}$$

where $\Psi(\lambda_j;\theta^1) = \mathrm{diag}\left(\lambda^\gamma, \lambda^\delta e^{-i(\pi-\lambda_j)(\delta-\gamma)/2}\right)$, the superscript asterisk denotes the conjugate transpose, $I_z(\lambda_j)$ is the tapered periodogram matrix of z_t and $\sum_{j(p,\kappa)}^{m}$ is the sum taken over $j = g, 2g, \ldots, m$ for $j \leq \lfloor \kappa m \rfloor$ and g the order of the taper. This estimator allows to capture the phase parameter (or time delay), modeled as $\tilde{\gamma} = (\delta - \gamma)\pi/2$, generally non-null in presence of spectra divergence at zero-frequency. Shimotsu (2012) also introduces a trimming parameter, κ, to control the behavior of the objective function when $\delta-\gamma > 1/2$. In the context of our application, this trimming parameter takes the following values $\kappa = \{0.05, 0.04\}$. The second step consists of an application of the 2S-ELW of Shimotsu (2010) to estimate θ^* as $\theta^* = \arg\min_{\theta\in\Theta} R^*(\theta)$ given the concentrated objective function,

$$R^*(\theta) = \log\det\tilde{\Omega}^*(\theta) - 2(\gamma+\delta)\frac{1}{m}\sum_{j=1}^{m}\log\lambda_j, \tag{3.9}$$

$$\tilde{\Omega}^*(\theta) = \frac{1}{m}\sum_{j=1}^{m}\mathrm{Re}\left[I_{\Delta^{\theta^1}z}(\lambda_j;\beta)\right], \qquad I_{\Delta^{\theta^1}z}(\lambda_j;\beta) = \omega_{\Delta^{\theta^1}z}(\lambda_j;\beta)\omega^*_{\Delta^{\theta^1}z}(\lambda_j;\beta) \tag{3.10}$$

$$\omega_{\Delta^{\theta^1}z}(\lambda_j;\beta) = \begin{pmatrix} \omega_{\Delta^\gamma(q_{t,a}-\beta q_{t,b})}(\lambda_j) \\ \omega_{\Delta^\delta q_{t,b}}(\lambda_j) \end{pmatrix} \tag{3.11}$$

with Θ the parameter space of θ^* and $\omega_z(\lambda_j)$ the discrete Fourier transform of z_t. Denoting $\hat{\theta}$ the estimate of θ from the aforementioned tapered LW estimator, the 2S-ELW-FC estimator of Shimotsu (2012) is defined as $\theta^* = \hat{\theta} - \left((\partial^2/\partial\theta\partial\theta')R^*(\hat{\theta})\right)^{-1}\left((\partial/\partial\theta)R^*(\hat{\theta})\right)$.[15] The author shows that $m^{1/2}\mathrm{diag}\left\{\lambda_m^{-v_0}, 1, 1\right\}(\theta^* - \theta_0) \xrightarrow{d} N(0, \Xi^{-1})$ as $n \to \infty$ when $v_0 \in (0, 1/2)$ and $m^{1/2}(\theta^{1*} - \theta_0^1) \xrightarrow{d} N(0, \Xi_{\theta^1}^{-1})$ while $(\beta^* - \beta_0) = O_p(n^{-v_0})$ as $n \to \infty$ when $v_0 \in (1/2, 3/2)$ with v_0 the true value of $v = (\delta - \gamma)$. Exploiting $\tilde{\Omega}^*(\theta)$, this procedure is also able to estimate the off-diagonal parameter, ρ, of the residuals covariance matrix (i.e. endogeneity parameter). The cointegration analysis is reported in Tables 3.7 to 3.11. Tables regroup the rank analysis and the regression-based approach.[16] In some cases, we apply a penalty parameter to β helps for convergence and improves the results.

3.4 Empirical results and discussion

As a first step, we employ the cointegration rank procedure to check for cointegration among the system composed of the five RERs, both in terms of US dollar and yen, before moving on to bivariate setting. Table 3.6 displays the results for each sample periods under consideration. Concerning the full sample, the results quite clearly indicate the presence of at least two but possibly three cointegrating

15. According to the author, an iterative procedure may result from estimation of θ^* without modifying the asymptotic distribution and may improve the finite sample properties.

16. We also estimate the model using the parametric quasi-maximum likelihood estimator of fractional cointegration (QML-FC) suggested by de Truchis (2013). This estimator is not originally designed to handle weak fractional cointegration although the author documents a Monte Carlo experiments that reveals good finite sample properties in this case. Consequently, our analysis does not rely on these results. Nonetheless, this study reveals that short-run dynamics of the cointegrating errors are fairly complex, and encourage us to select a small bandwidth (i.e. $m = n^{0.6}$) to avoid the contamination of the low frequencies from the higher frequencies that are dominated by short-run dynamics.

relations in terms of US dollar. Indeed, for the case with the largest $\upsilon(n)$, i.e. $\upsilon(n) = m_1^{-0,25}$, we find evidence that the rank may be as high as three. Concerning the yen, the results indicate three cointegrated relations, regardless of $\upsilon(n)$. Interestingly, the cointegration rank analysis displays one cointegrating relation in terms of US dollar but three in terms of yen in the pre-crisis period and these results are robust to the choice of $\upsilon(n)$. Finally, we find two cointegrating relations in terms of US dollar and four in terms of yen in the post-crisis period. Comparing this with the results of the pre-crisis period, we reach two important preliminary conclusions. First, the Southeast Asian RERs appear to be more closely tied after the Asian financial crisis since we find for the two numeraires one additional cointegrating relation in the post-crisis period. Second, our results suggest that the yen is better suited than the US dollar as a monetary anchor for the region. Indeed, for each sample the number of cointegrating relations is higher in terms of yen.

Table 3.6: Rank estimates among the five Southeast Asian RERs using the model selection procedure with $\hat{P}(\bar{\delta}_*)$

	1975:01 - 2011:12 (n = 444)		1975:01 - 1997:05 (n = 269)		1998:06 - 2011:12 (n = 162)	
	$\upsilon(n) = m_1^{-0,35}$	$\upsilon(n) = m_1^{-0,25}$	$\upsilon(n) = m_1^{-0,35}$	$\upsilon(n) = m_1^{-0,25}$	$\upsilon(n) = m_1^{-0,35}$	$\upsilon(n) = m_1^{-0,25}$
$L(u)$ (USD)						
$L(1)$	-3.442	-2.826	-3.277	-2.664	-3.062	-2.459
$L(2)$	-3.565	-3.072	-3.354	-2.863	-3.198	-2.716
$L(3)$	-3.568	-3.199	-3.152	-2.784	-3.282	-2.920
$L(4)$	-3.461	-3.214	-2.602	-2.357	-3.065	-2.824
$L(5)$	-3.080	-2.956	-1.713	-1.590	-2.772	-2.652
$\hat{r}$	2	3	1	1	2	2
$L(u)$ (YEN)						
$L(1)$	-3.442	-2.826	-3.277	-2.664	-3.062	-2.459
$L(2)$	-3.673	-3.180	-3.567	-3.077	-3.411	-2.929
$L(3)$	-3.854	-3.484	-3.814	-3.446	-3.719	-3.357
$L(4)$	-3.985	-3.739	-3.995	-3.750	-3.948	-3.707
$L(5)$	-3.640	-3.517	-3.802	-3.679	-3.990	-3.869
$\hat{r}$	3	3	3	3	4	4

We now turn to the examination of the pairwise relationships, by applying both the cointegration test and the semi-parametric regression-based approach. The empirical study conducted in this section is based on the examination of fractional cointegration relationships for 20 pairs of Southeast Asian RERs. We consider two numeraires: USD and YEN. As emphasized by Sun and Simons (2011), each potential member decides on a unilateral basis whether to join a CCA. Accordingly, bivariate models are more likely to deliver straightforward evidence for whether one country shares co-movement with another potential member. The results are presented, by country, in Table 3.7 till Table 3.11 for the full sample and the two sub-samples. We discuss estimation results obtained from the full sample and conclude with the sub-samples analysis.

Table 3.7: Cointegration analysis of Malaysia

	1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
USD	Indo.	Phil.	Sing.	Thai.	Indo.	Phil.	Sing.	Thai.	Indo.	Phil.	Sing.	Thai.
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	0	1	1	0	0	1	0	0	0	1	0
$m_1^{-0.25}$	1	1	1	1	0	0	1	0	0	1	1	0
$m_1^{-0.15}$	1	1	1	1	0	0	1	0	1	1	1	1
						Regression analysis						
$\hat{\beta}$	0.285	1.129	1.262	1.204	0.372	0.750	0.447	1.279	0.297	0.427	0.703	0.503
	(0.071)	(0.573)	(0.635)	(0.222)	(0.295)	(0.412)	(2.113)	(0.904)	(0.392)	(0.156)	(0.405)	(0.180)
$\hat{\gamma}$	0.685	0.997	0.986	0.913	0.870	0.847	1.031	0.828	1.016	0.940	1.072	0.802
	(0.080)	(0.072)	(0.080)	(0.069)	(0.084)	(0.085)	(0.091)	(0.084)	(0.103)	(0.105)	(0.107)	(0.107)
$\hat{\delta}$	0.923	1.077	1.082	1.071	0.990	0.978	1.003	0.918	1.076	1.164	1.185	0.998
	(0.080)	(0.072)	(0.080)	(0.069)	(0.084)	(0.085)	(0.091)	(0.084)	(0.103)	(0.105)	(0.107)	(0.107)
ρ	-0.176	-0.648	-0.280	-0.739	-0.648	-0.616	0.397	-0.664	-0.580	-0.458	-0.405	-0.395
YEN												
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	1	1	1	0	1	1	1	1	1	1	1
$m_1^{-0.25}$	1	1	1	1	1	1	1	1	1	1	1	1
$m_1^{-0.15}$	1	1	1	1	1	1	1	1	1	1	1	1
						Regression analysis						
$\hat{\beta}$	0.581	1.240	1.609	1.202	0.598	1.163	1.437	1.209	0.385	0.527	0.979	0.624
	(0.532)	(2.519)	(1.025)	(0.387)	(0.791)	(0.937)	(1.403)	(0.325)	(0.426)	(0.131)	(0.204)	(0.173)
$\hat{\gamma}$	0.972	0.983	1.025	0.915	1.072	1.027	1.081	0.876	1.060	0.675	0.653	0.758
	(0.077)	(0.068)	(0.071)	(0.072)	(0.093)	(0.078)	(0.082)	(0.087)	(0.111)	(0.105)	(0.104)	(0.106)
$\hat{\delta}$	0.909	0.995	1.062	0.989	1.132	1.061	1.104	0.994	0.959	0.937	0.814	0.980
	(0.077)	(0.068)	(0.071)	(0.072)	(0.093)	(0.078)	(0.082)	(0.087)	(0.111)	(0.105)	(0.104)	(0.106)
ρ	-0.488	-0.787	-0.716	-0.681	-0.234	-0.808	-0.738	-0.539	0.169	0.463	0.504	-0.350

Notes: Standard deviations are displayed in parentheses (.).

Table 3.8: Cointegration analysis of Thailand

	1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
USD	Indo.	Mal.	Phil.	Sing.	Indo.	Mal.	Phil.	Sing.	Indo.	Mal.	Phil.	Sing.
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	1	0	0	0	0	0	0	1	0	1	0
$m_1^{-0.25}$	1	1	1	1	0	0	0	0	1	0	1	1
$m_1^{-0.15}$	1	1	1	1	0	0	0	0	1	1	1	1
						Regression analysis						
$\hat{\beta}$	0.351	0.716	0.985	0.566	0.210	0.537	0.611	0.597	0.615	1.679	0.811	1.235
	(0.724)	(0.168)	(0.292)	(0.243)	(0.708)	(4.813)	(0.224)	(0.448)	(0.068)	(0.288)	(0.054)	(0.53)
$\hat{\gamma}$	0.907	0.806	0.885	0.810	0.959	0.884	0.934	0.876	0.670	0.775	0.744	0.972
	(0.08)	(0.08)	(0.077)	(0.08)	(0.093)	(0.093)	(0.078)	(0.092)	(0.097)	(0.106)	(0.107)	(0.108)
$\hat{\delta}$	0.873	1.055	1.059	1.102	1.002	0.897	1.127	1.017	1.041	1.121	1.208	1.140
	(0.08)	(0.08)	(0.077)	(0.08)	(0.093)	(0.093)	(0.078)	(0.092)	(0.097)	(0.106)	(0.107)	(0.108)
ρ	-0.292	0.256	-0.425	0.199	-0.275	-0.303	-0.791	-0.337	-0.651	-0.350	-0.371	-0.355
YEN												
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	1	1	1	1	0	1	1	1	1	1	1	1
$m_1^{-0.25}$	1	1	1	1	1	1	1	1	1	1	1	1
$m_1^{-0.15}$	1	1	1	1	1	1	1	1	1	1	1	1
						Regression analysis						
$\hat{\beta}$	0.229	0.789	1.044	0.938	0.480	0.797	0.961	1.193	0.622	1.226	0.860	1.407
	(0.021)	(0.105)	(0.261)	(0.103)	(0.061)	(0.115)	(0.18)	(0.249)	(0.263)	(3.175)	(0.156)	(0.285)
$\hat{\gamma}$	0.695	0.809	0.776	0.814	0.838	0.844	0.840	0.923	0.747	0.914	0.748	0.734
	(0.075)	(0.079)	(0.076)	(0.081)	(0.093)	(0.094)	(0.084)	(0.089)	(0.111)	(0.103)	(0.111)	(0.107)
$\hat{\delta}$	1.168	1.066	0.897	1.156	1.241	1.108	1.013	1.079	0.879	0.928	0.960	0.911
	(0.075)	(0.079)	(0.076)	(0.081)	(0.093)	(0.094)	(0.084)	(0.089)	(0.111)	(0.103)	(0.111)	(0.107)
ρ	0.453	0.328	-0.506	0.069	0.242	0.066	-0.644	-0.466	-0.213	-0.585	-0.195	-0.418

Notes: Standard deviations are displayed in parentheses (.).

Table 3.9: Cointegration analysis of Philippines

	1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
USD	Indo.	Mal.	Sing.	Thai.	Indo.	Mal.	Sing.	Thai.	Indo.	Mal.	Sing.	Thai.
$\upsilon(n)$					Rank Analysis							
$m_1^{-0.35}$	0	0	0	0	0	0	0	0	0	0	0	1
$m_1^{-0.25}$	1	1	0	1	0	0	0	0	1	1	0	1
$m_1^{-0.15}$	1	1	1	1	0	0	0	0	1	1	1	1
					Regression analysis							
$\hat{\beta}$	0.264	0.542	1.347	0.796	0.209	0.542	0.932	1.051	0.612	1.957	1.440	1.113
	(0.457)	(3.129)	(0.854)	(0.195)	(6.314)	(2.292)	(1.295)	(1.573)	(0.402)	(0.834)	(0.814)	(0.427)
$\hat{\gamma}$	1.005	1.036	0.959	0.873	0.991	0.959	0.950	1.091	0.958	1.078	0.993	0.912
	(0.081)	(0.081)	(0.078)	(0.079)	(0.094)	(0.094)	(0.09)	(0.083)	(0.1)	(0.103)	(0.103)	(0.104)
$\hat{\delta}$	0.953	1.018	1.063	1.098	0.998	0.918	1.028	1.013	1.074	1.257	1.128	1.076
	(0.081)	(0.081)	(0.078)	(0.079)	(0.094)	(0.094)	(0.09)	(0.083)	(0.1)	(0.103)	(0.103)	(0.104)
ρ	-0.144	0.170	-0.420	-0.322	-0.223	-0.192	-0.454	-0.725	-0.647	-0.531	-0.558	-0.514
YEN												
$\upsilon(n)$					Rank Analysis							
$m_1^{-0.35}$	0	1	1	1	1	1	1	1	1	1	1	1
$m_1^{-0.25}$	1	1	1	1	1	1	1	1	1	1	1	1
$m_1^{-0.15}$	1	1	1	1	1	1	1	1	1	1	1	1
					Regression analysis							
$\hat{\beta}$	0.399	0.677	1.286	0.869	0.469	0.771	1.188	0.965	0.591	1.250	1.493	1.038
	(0.104)	(0.844)	(0.303)	(0.453)	(0.05)	(0.418)	(0.256)	(0.178)	(0.314)	(0.617)	(0.586)	(0.144)
$\hat{\gamma}$	0.788	0.962	1.006	0.831	0.699	0.945	1.000	0.733	1.145	0.968	0.883	0.718
	(0.081)	(0.076)	(0.081)	(0.081)	(0.089)	(0.089)	(0.094)	(0.094)	(0.111)	(0.108)	(0.108)	(0.111)
$\hat{\delta}$	1.070	1.008	1.154	0.914	1.155	1.062	1.183	1.001	0.974	0.880	0.986	0.992
	(0.081)	(0.076)	(0.081)	(0.081)	(0.089)	(0.089)	(0.094)	(0.094)	(0.111)	(0.108)	(0.108)	(0.111)
ρ	0.087	0.533	-0.092	0.068	0.424	0.482	0.107	0.168	-0.171	-0.408	-0.356	-0.138

Notes: Standard deviations are displayed in parentheses (.).

Table 3.10: Cointegration analysis of Indonesia

	1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
USD	Mal.	Phil.	Sing.	Thai.	Mal.	Phil.	Sing.	Thai.	Mal.	Phil.	Sing.	Thai.
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	0	0	0	0	0	0	0	0	0	0	1
$m_1^{-0.25}$	1	1	0	1	0	0	0	0	0	0	1	1
$m_1^{-0.15}$	1	1	0	1	0	0	0	0	1	1	1	1
						Regression analysis						
$\hat{\beta}$	1.835	2.099	2.158	2.256	2.332	1.811	0.828	3.136	2.077	0.933	1.477	1.288
	(0.237)	(0.475)	(2.512)	(0.2)	(1.779)	(1.214)	(1.556)	(6.028)	(0.369)	(0.328)	(1.071)	(0.276)
$\hat{\gamma}$	0.663	0.893	1.007	0.810	1.029	0.896	1.100	0.991	0.899	0.850	0.967	0.732
	(0.08)	(0.077)	(0.08)	(0.072)	(0.081)	(0.081)	(0.092)	(0.081)	(0.111)	(0.111)	(0.112)	(0.112)
$\hat{\delta}$	1.039	1.124	1.092	1.171	0.926	0.994	0.990	0.959	1.319	1.181	1.130	1.075
	(0.08)	(0.077)	(0.08)	(0.072)	(0.081)	(0.081)	(0.092)	(0.081)	(0.111)	(0.111)	(0.112)	(0.112)
ρ	-0.176	-0.450	-0.249	-0.609	-0.777	-0.725	-0.352	-0.764	-0.143	-0.180	-0.074	-0.067
YEN												
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	0	0	1	0	1	0	0	1	1	1	1
$m_1^{-0.25}$	1	1	0	1	1	1	1	1	1	1	1	1
$m_1^{-0.15}$	1	1	0	1	1	1	1	1	1	1	1	1
						Regression analysis						
$\hat{\beta}$	1.596	2.004	2.531	1.955	1.555	1.840	2.211	1.892	1.675	1.085	1.930	1.378
	(0.141)	(0.357)	(0.562)	(0.166)	(0.303)	(0.301)	(0.496)	(0.316)	(0.715)	(0.494)	(1.238)	(0.339)
$\hat{\gamma}$	0.764	0.817	0.972	0.689	0.945	0.843	0.988	0.888	0.832	0.804	0.893	0.726
	(0.074)	(0.068)	(0.069)	(0.068)	(0.082)	(0.074)	(0.077)	(0.078)	(0.103)	(0.11)	(0.107)	(0.11)
$\hat{\delta}$	1.145	1.006	1.165	1.011	1.167	1.007	1.148	1.088	0.953	0.957	0.973	0.934
	(0.074)	(0.068)	(0.069)	(0.068)	(0.082)	(0.074)	(0.077)	(0.078)	(0.103)	(0.11)	(0.107)	(0.11)
ρ	-0.536	-0.754	-0.729	-0.733	-0.677	-0.883	-0.813	-0.778	-0.538	-0.228	-0.401	-0.261

Notes: Standard deviations are displayed in parentheses (.).

Table 3.11: Cointegration analysis of Singapore

	1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
USD	Indo.	Mal.	Phil.	Thai.	Indo.	Mal.	Phil.	Thai.	Indo.	Mal.	Phil.	Thai.
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	1	0	0	0	1	0	0	0	1	0	1
$m_1^{-0.25}$	0	1	0	1	0	1	0	0	1	1	0	1
$m_1^{-0.15}$	0	1	1	1	0	1	0	0	1	1	1	1
						Regression analysis						
$\hat{\beta}$	0.099	0.221	0.490	0.374	0.051	0.174	0.503	0.554	0.394	1.311	0.586	0.690
	(0.134)	(0.573)	(0.192)	(0.201)	(0.457)	(1.501)	(1.443)	(2.205)	(0.296)	(0.217)	(0.117)	(0.236)
$\hat{\gamma}$	1.087	1.071	0.990	0.948	1.028	1.000	1.005	0.957	0.981	0.912	0.838	0.893
	(0.081)	(0.078)	(0.078)	(0.08)	(0.094)	(0.086)	(0.089)	(0.093)	(0.104)	(0.108)	(0.104)	(0.106)
$\hat{\delta}$	0.965	1.010	1.162	1.101	0.955	1.033	1.042	0.990	1.082	1.230	1.179	1.086
	(0.081)	(0.078)	(0.078)	(0.08)	(0.094)	(0.086)	(0.089)	(0.093)	(0.104)	(0.108)	(0.104)	(0.106)
ρ	-0.026	0.465	-0.400	-0.213	-0.061	0.600	-0.484	-0.270	-0.534	-0.357	-0.463	-0.421
YEN												
$\upsilon(n)$						Rank Analysis						
$m_1^{-0.35}$	0	1	1	1	0	1	1	1	1	1	1	1
$m_1^{-0.25}$	0	1	1	1	1	1	1	1	1	1	1	1
$m_1^{-0.15}$	0	1	1	1	1	1	1	1	1	1	1	1
						Regression analysis						
$\hat{\beta}$	0.268	0.468	0.685	0.602	0.365	0.616	0.769	0.776	0.404	0.892	0.574	0.653
	(0.309)	(1.04)	(0.24)	(0.138)	(0.165)	(2.882)	(1.458)	(0.386)	(0.084)	(0.291)	(0.161)	(0.146)
$\hat{\gamma}$	1.111	1.031	1.106	1.150	0.999	1.092	1.091	0.954	1.370	0.876	0.754	0.771
	(0.08)	(0.075)	(0.08)	(0.081)	(0.094)	(0.087)	(0.09)	(0.094)	(0.112)	(0.107)	(0.112)	(0.112)
$\hat{\delta}$	1.014	1.005	1.005	0.927	1.224	1.083	1.073	1.043	0.955	0.798	0.925	0.975
	(0.08)	(0.075)	(0.08)	(0.081)	(0.094)	(0.087)	(0.09)	(0.094)	(0.112)	(0.107)	(0.112)	(0.112)
ρ	-0.233	0.590	-0.263	-0.030	0.164	0.598	-0.457	0.055	-0.049	-0.467	0.086	0.066

Notes: Standard deviations are displayed in parentheses (.).

The cointegration rank reveals several pairs of RERs which are fractionally cointegrated. A closer reading reveals some important dissimilarities depending on the base currency. The majority of the Southeast Asian countries are mutually fractionally cointegrated when the base currency is the yen, suggesting that further monetary integration is possible. However, the possibility of a core-group composed of the five Southeast Asian countries should be excluded as Indonesia and Singapore do not share a common long-run equilibrium over the full period. Regarding the US dollar, we find mixed results since the rank estimates is quite sensitive to the $\upsilon(n)$ parameter in contrast to the yen-based estimates. Nonetheless, we find some clusters as potential candidates such as Thailand, Malaysia and Singapore.

The regression-based approach concurs with the rank analysis, especially for Indonesia, The Philippines and Thailand. Indeed, the estimates of β are most often significant, while the fractional integration order of the residuals (i.e. $\hat{\gamma}$) are lesser than the fractional integration order of the regressors (i.e. $\hat{\delta}$) for several pairwise RERs, both in terms of US dollar and yen. However, the evidence concerning Malaysia and Singapore do not corroborate the results obtained from the rank analysis. In such a case, we must be careful when interpreting the estimation results. However, differences in macroeconomic fundamentals over this period, such as demography, specialization patterns or the initial level of output per capita, can justify lesser integration of Singapore's economy with that of its Southeast Asian neighbors. At this stage, such deep heterogeneity should restrict the possibility of a monetary union in Southeast Asia involving Singapore.

To examine the strength of the cointegration relationships, we analyze the gap between the integration order of the regressor ($\hat{\delta}$) and the integration order of the residuals ($\hat{\gamma}$). Weak fractional cointegration relationships are found in all cases (i.e. $\hat{\delta} - \hat{\gamma} < 1/2$ with $\hat{\delta} > 1/2$). Accordingly, deviations are highly persistent but mean reversion occurs in the long run. This finding has important implications for the region. First, fundamentals driving RERs in Southeast Asia share a common equilibrium in the long run. Second, RERs are tied together through a long memory process which contrast with results of all previous studies. As a result of this finding, empirical models of G-PPP theory that ignore such a feature should be misspecified.

The sign and the value of the long-run β coefficient provide information for examining the way countries adjust to shocks. We find positive long-run coefficient in all cases, thus implying symmetric adjustment and exposure to macroeconomic disturbances (e.g. global shocks in particular industries, such as textiles and electronic sector, monetary policy shocks in the anchor country). This is another important feature pleading in favor of further monetary integration among those countries. Long-run elasticities are particularly high between the RERs of Indonesia and other countries. For instance, the yen-based (resp. the US dollar-based) RER of Indonesia has depreciated from 1.596% (resp. 1,835%), 2.004% (resp. 2.099%), 2.531% (resp. 2.158) and 1.955% (resp. 2.256%) following a 1% depreciation of the Malaysian ringgit, the Philippine peso, the Singaporean dollar and the Thai bath, respectively. These results may be explained by the sharp depreciation of the Indonesian rupiah during the Asian crisis.[17] Overall, with regard to the difference between the two numeraires, long-run coefficients are higher in the case of the yen.

In summary, we can state that neither a yen bloc nor a US dollar bloc has formed in Southeast Asia and a monetary union involving all ASEAN-5 countries is actually compromised. On the whole, these findings are consistent with some previous researches that focused on G-PPP as a precondition for a CCA

17. Indonesia and Thailand were the countries most affected by the crisis. Except for Singapore which has weathered the crisis better than most Asian economies, the currencies of the ASEAN-5 countries had depreciated 20-30% on average, while Indonesian rupiah had depreciated more than 50%.

(Sun and Simons , 2011; Wilson and Choy , 2007). In the perspective of a future CCA, the results suggest that unilateral US dollar-based currency union seems less suited than for the basket alternative, which is consistent with a number of previous studies that emphasize the need to promote a common currency basket with a large weight for the yen (Mishra and Sharma , 2010; Ogawa and Kawasaki , 2008). As a result, greater stabilization against the yen should be desirable given the economic interdependence between Japan and its Southeast Asian partners. [18]

Following the structural change analysis conducted in Section 3.3, we compare estimation results of the two sub-samples (1975:1-1997-5 for the pre-crisis period and 1998:6-2011:12 for the post-crisis period) considering the substantial impact of the 1997-98 crisis. One major consequence of the crisis has been the emergence of a higher degree of policy cooperation in trade and monetary areas. According to Sato and Zhang (2006), economic regionalism in Southeast Asia has resulted in greater interdependence of economic structures and contributed to the co-movements of real output variables in the long run. Therefore, we must consider how the Asian financial crisis has influenced macroeconomic factors that guide the RERs within the region.

In terms of US dollar, the rank analysis displays only one cointegration relationship, i.e. between Singapore and Malaysia, in the pre-crisis sample. However, the estimate of β is not significant while the gap between $\hat{\gamma}$ and $\hat{\delta}$ is close to zero. Overall, there is no reliable evidence either on the rank or regression analyses suggesting that US dollar-based RER of the Southeast Asian countries are tied together before the Asian financial crisis. In sharp contrast, the rank analysis displays several cointegration relationships among yen-based RERs although the results are relatively sensitive to the tunning parameter $\upsilon(n)$ in the case of Indonesia. In addition, the estimates of β are not significant for some pairwise RERs such as Malaysia with Indonesia, the Philippines, Singapore and Singapore with Malaysia and the Philippines. However, we find that the integration order of residuals is less than the integration order of regressor confirming, in a few cases, the rank estimates. For instance, the strength of the cointegration relationship is relatively much stronger for pairwise RERs including Indonesia, the Philippines and Thailand. Comparing these results with those of the post-crisis sample, we observe that the Southeast Asian RERs are better tied together after the crisis. Indeed, the rank estimates indicate some fractionally cointegrated US dollar-based RERs for which the strength of the relationship is relatively strong. This finding is particularly significant, again, for Indonesia, Thailand and the Philippines while in other cases, the regression analysis do not support the rank estimates. As such, the results have to be received with reservations. Finally, we find strong evidence in favor of fractional cointegration among yen-based RERs after the Asian crisis. More importantly, the robustness of these results to different tunning parameter $\upsilon(n)$ is always validated contrary to the results of the previous samples. At the same time, this concurs most often with the regression analysis for which we observe significant estimates of β and, sometimes, non-negligible $I(\hat{\gamma} - \hat{\delta})$ values. Furthermore, the results indicate that equilibrium errors respond faster to shocks in the post-crisis period, thus implying less persistent deviations toward the long-run equilibrium. Accordingly, the real income process underlying the evolution of the RERs has been more interrelated after the crisis, as a result of exchange rate policy reforms although evidences

18. First, this argument relates to the significant role played by Japan and the yen into the region in terms of external trade relation and financial transactions (see Rajan , 2002). Second, the de facto US dollar-peg adopted by most Southeast Asian countries before the crisis is believed to be some of the main causes of the crash of 1997-98. After 1985 the deprecation of the US dollar led to the depreciation of Southeast Asian currencies against the yen. However, the nearly 50% appreciation of US dollar between June 1995 to April 1997 led to overvaluation of regional RERs that worsened the terms of trade and caused massive capital outflows. It is commonly assumed that if they had given greater weight to the yen there would have been lower degrees of exchange rate misalignments.

are not sufficient enough to assert the existence of an OCA over the last decade.

The contrast between the two periods can be explained as follows: before the crisis, shifts in bilateral RERs in terms of the US dollar have taken place mainly through inflation differentials since nominal exchange rates was anchored to the US dollar.[19] Such finding suggests structural differences across countries over this period such as non-uniform productivity growth rates, heterogeneous structural policies and asymmetric resilience to shocks. Since the crisis, Asian exchange rate policies have moved toward greater flexibility while monetary policies are conducted under inflation targeting frameworks. This strategy appears to have been successful in enabling those countries to maintain low inflation rates, something that they were unable to do in the past, conducive to a balanced and sustainable export-led growth. Southeast Asian currencies remain fairly managed, against the US dollar or a currency basket, in order to maintain a stable NEER and price competitiveness. Thus, cointegrating relationships found in the post-crisis period should reflect, to some extent, implicit monetary policy coordination and similarity of trade structure. Indeed, the increase of intra-regional trade, from 30.72% to 45.92% since 1975, is the result of an intense intra-industry trade networking within the region in which Multinational Corporation and their affiliates exchange parts and components, intermediate inputs and finished manufactured products. It is quite likely that such a pattern has tended to synchronize economic fluctuations among Southeast Asian countries.

3.5 Conclusion

This paper analyzed the feasibility of a monetary union for the ASEAN-5. The analysis covers a period from January 1975 to December 2011 and two sub-periods to account for the effects of the Asian financial crisis. We applied G-PPP theory to examine long-term co-movements among the RERs. To this end, a new estimator of fractional cointegration was introduced. In contrast to all previous studies, our model takes into account a wider range of dynamic reversion toward the long-run equilibrium. Our results shown that long-term stability is fragile because the strength of the cointegration relationship is weak. Nonetheless, we found that the equilibrium errors respond more quickly to shocks in the post-crisis sample resulting in less persistent deviations toward the common equilibrium. This is an encouraging result when we consider the desirability of a currency union in the future.

The full sample analysis highlights different sub-groups as potential candidates: Indonesia, Malaysia, the Philippines, and Singapore as well as Indonesia, Malaysia, the Philippines and Thailand in terms of Japanese yen and Thailand, Malaysia and Singapore in terms of US dollar. The results shown that an exchange rate policy based on a currency basket is probably more effective in ensuring the regional stability of the RERs. Actually, most of the Southeast Asian countries operate under managed float regimes by controlling their margins of fluctuations. However, there is not yet a formal form of exchange rates coordination that may lead to strengthen intra-regional exchange rate stability and which could then establish a wide zone of relative price stability. Our results support the case for further monetary integration in the ASEAN-5 and call for a set of ambitious reforms that aim to increase intra-regional trade and real linkages, such as informal forms of exchange rates policy coordination where the yen and the US dollar play an essential role. As intra-regional trade and FDI flows continue to increase, the region will become more interdependent, thus paving the way for a full-fledged monetary union.

19. That is true to a lesser extent for Singapore. Officially, the monetary authority has been targeting its nominal effective exchange rate (NEER) since 1981.

Chapter

4

Analyzing Financial Integration in Asia Through Fractional Cointegration of Stock Market Volatility

4.1 Introduction

Financial integration issue is closely related to cross-country links in stock markets and more precisely to the correlation of returns and volatility (see, e.g., Beckaert and Harvey , 1995). Given that volatility of stock markets reflects to a large extent the risk exposure of assets, we expect that existence of cross-market correlations should reveal common risk premium associated with an identical risk exposure. Accordingly, integrated stock markets are more likely to display greater correlation among their volatility processes because the source of risk is the same, while the opposite is true for segmented markets. Harvey (1995) argues that stock markets in emerging countries are more likely to be influenced by local events, suggesting a lower degree of financial integration, while developed markets are more likely to be tied through a global integration process because the risk exposures are likely influenced by global variables (see, also, Engle and Susmel , 1993). This issue has been widely investigated in the literature since many financial and regulatory activities (e.g. pricing of financial instruments, performance evaluation, risk hedging strategies and portfolio allocation) depend upon the perceived commonality in volatility movements. In this regard, understanding interrelations among closely related stock markets is crucial.

In this paper, we explore the presence of common stochastic trends among integrated volatilities of several Asian equity markets (Japan, Hong Kong, Singapore, Malaysia, Thailand, Indonesia, the Philippines) to assess the degree of financial integration in this area. We support that existence of long-run relationship between their integrated volatilities should reveal at least a partial financial integration while absence of commonalities should indicate segmented markets. Indeed, cross-border correlation in volatilities could exist throughout a common long-run component although the transitory components might temporary diverge. In such cases, stock markets are cointegrated, implying that the volatilities converge in the long run and respond in similar ways to shocks. This entails that volatilities are driven in the long run by a common information arrival process (i.e. the information events causing the volatility are the same) that should reflect commonalities in the market fundamentals (e.g. trade connections, similar economic and institutional structures and cultures). Considering the well documented influence of the US on Asian stock markets (see, e.g., Ng , 2000; Chuang et al. , 2007), we also include in our panel the Standard & Poor's 500 to test whether volatilities in these markets are cointegrated with the US stock market.

Previous empirical studies rely on numerous econometric approaches including among other, VAR and cointegration analysis, GARCH class of models or common-component approach (see, e.g., Yu et al. , 2010, for a survey of high-frequency integration measures). Since higher correlation between second moment of returns generally implies greater integration among stock markets and higher co-movement, constant and dynamic correlation analysis of conditional variance has been widely used to capture the extent to which the covariances with world or regional markets are relevant. For instance, Miyakoshi (2003) shows that the volatility of the Asian financial markets is influenced more by the Japanese market than by the US. Tai (2007) examines whether Asian emerging stock markets have become integrated into world capital markets and find full market integration after the official liberalization date,[1] while Chambert and Gibson (2008) find that Asian countries still remain segmented, even if the level of stock market integration has been trending upwards more recently. Yu et al. (2010) analyze stock markets correlation using dynamic conditional correlation model and find that all stock markets in Asia show increasing correlation with each other in the past year (see, also, Ng , 2000; Worthington and Higgs , 2004; Click and Plummer , 2005; Chuang et al. (2007); Beirne et al. , 2010; Abbas et al. , 2013; Lee , 2013).

However, all aforementioned studies focus on short term correlation and, therefore, the transitory component of the variances and covariances. Accordingly, they neglect the persistent nature of volatility which is of importance regarding the possibility of long-run correlation. Our methodology employs recent fractional cointegration techniques and relies on the rank analysis as well as the regression-based analysis, accommodating in both cases the possible non-stationarity of the volatility processes. Considering that volatilities are persistent, the fractional cointegration is a convenient approach to capture the share of the volatility in both markets which has a common origin. The pioneer work of Bollerslev and Engle (1993) addresses this issue by extending the cointegration theory to conditional variance (they term this concept co-persistence). Li (2012) deals with some drawbacks of Bollerslev and Engle (1993) and find strong evidences of financial integration in European markets. However, the multivariate integrated GARCH framework of Bollerslev and Engle (1993) and Li (2012) implies a very strong hypothesis: the volatility process possesses a unit root.[2] Brunetti and Gilbert (2000) relax this assumption, considering long memory in conditional variance.[3] Ho and Zang (2012) study the financial integration of the Greater China through a varying-correlation fractionally integrated asymmetric power GARCH model. They extensively discuss the co-persistence issue although they do not test for the presence of common stochastic trends. Dao and Wolters (2008) investigate co-persistence through multivariate stochastic volatility model but assuming a unit root in volatility processes. They find evidence of common stochastic volatility trends in several developed East Asian stock markets. In a pure theoretical study, da Silva and Robinson (2008) relax the unit root assumption of Dao and Wolters (2008) and develop a bivariate long memory stochastic volatility model. Concerning integrated variance, the pioneer work of Andersen et al. (2003) provides an heuristic analysis of fractional cointegration among the volatility of three exchange rates and concludes in favor of no cointegration. In a recent study, Cassola and Morana (2010) study the common sources of volatility in the euro Area money market interest rates.

1. The author uses the official liberalization dates reported by the International Finance Corporation.

2. The presence of a unit root in the volatility stochastic process seems too extreme inasmuch as it implies permanent shifts to long-term volatility forecasts, which is theoretically implausible.

3. See Figuerola-Ferretti and Gilbert (2008) for an application of the bivariate error correction mechanism FIGARCH model of Brunetti and Gilbert (2000). See also Bollerslev and Mikkelsen (1996) and Andersen et al. (2003) for a discussion about the long memory feature of the conditional and integrated variance.

In line with these studies, we investigate the presence of common stochastic trends among the integrated volatilities of several Asian stock markets. The extent to which emerging and developed markets in Asia are fractionally cointegrated is of importance for three reasons. First, the lack of commonality in the volatility processes (i.e. the absence of a common underlying trend) can emphasize the potential benefits of portfolio diversification, especially for investors who engage in volatility arbitrage on these markets. Second, intensified financial linkages in a world of high capital mobility may also generate the risk of volatility spillovers among closely related markets, but also undesirable macroeconomic effects (e.g. monetary expansion, inflationary pressures, real exchange rate appreciation). Accordingly, if volatilities respond in similar ways to shocks, policymakers and regulators within the region may find an interest to achieve greater degree of financial cooperation to enhance regional financial stability.[4] Moreover, financial integration is theoretically related to the possibility of sharing country-specific output shocks, both in "good" and "bad" times, via the cross-ownership of equities. Therefore, financial integration would potentially benefit the region through consumption smoothing and risk sharing, facilitating the adjustment in the face of asymmetric shocks. From the perspective of Asian monetary integration, this would make the adoption of a common currency more desirable because financial market integration would act as an insurance mechanism, which is of great importance for the smooth functioning of a monetary union (see, e.g, de Grauwe , 2012).[5]

To anticipate our main conclusions, the preliminary analysis reveals that, in most cases, the pairwise volatility processes exhibit a common degree of persistence. We also find strong evidences of co-persistence between volatilities of developed markets, confirming that Hong Kong and Singapore enter into a global integration process. Conversely, we find only little evidences of cointegration in variance among Malaysia, the Philippines, Indonesia and Thailand equity markets, confirming that emerging markets remain segmented.

The remainder of the paper is laid out as follows. The Section 4.2 details the dataset and the strategy of estimation. The Section 4.3 discuss the results and their economic implications. Section 4.4 concludes the paper.

4.2 Fractional cointegration estimation and testing

4.2.1 The data

Over the last three decades, the Asian economies have emerged as a pole of economic growth. The opening up of additional channels for cross-border linkages has contributed to more interrelated economies, which in turn, has given rise to greater co-movements with world stock markets. For instance, Park and lee (2011) indicate that the Asian countries have attracted about 57% of total financial inflows to emerging market economies over the last two decades, as a result of financial deregulation and capital account liberalization. Chinn and Ito (2008) construct a *de jure* index of capital account openness, also called *KAOPEN*, and find that the Asian countries have reached a high level of financial

4. The Asian governments have already embarked on several initiatives for regional financial cooperation in the aftermath of the 1997-98 crisis, including regional economic surveillance process (the Economic Review and Policy Dialog in May 2000), a liquidity support arrangement under the form of bilateral swaps (the Chiang Mai Initiative Multilateralized in 2009) and Asian bond markets developments (the Asian Bond Fund initiative of the Executive Meeting of East Asia Pacific Central Banks in 2003) However, it is generally argued that financial integration lags behind real integration, despite the political support toward greater financial cooperation (see, e.g., Kim and Lee , 2012).

5. The other analytical benefits refer to the positive impact of capital flows on domestic investment and growth, the additional discipline on macroeconomic policies and the efficiency improvements of the banking sector (Agénor , 2003).

openness compared to other groups of emerging countries, although the rate of financial opening slowed down in the aftermath of the 1997-98 crisis.

We concentrate our analysis on the national stock market index of seven Asian countries, including both developed and emerging markets: the Stock Exchange of Thailand (SET), the Straits Time Index of Singapore (STI), the Philippines Stock Exchange Index (PSEI), the Kuala Lumpur Composite Index (KLCI, Malaysia), the Jakarta Composite Index (JCI, Indonesia), the Hang Seng Index (HSI, Hong Kong) and the Nikkei 225 (NKY, Japan). Our data set cover the period January 1, 2000 to November 29, 2012, and the sample size is 3315.[6] Three of these markets are regarded as developed (Hong Kong, Japan and Singapore) while others are categorized as emerging (Indonesia, Malaysia, the Philippines and Thailand). Tokyo and Hong Kong are the two largest international financial centers in Asia in terms of market capitalization. Moreover, these two countries with Singapore are the more financially open economies in Asia according to the *KAOPEN* index of Chinn and Ito (2008) (see Table 4.1).[7] Finally, according to Click and Plummer (2005), the five most advanced members of the ASEAN (i.e. Singapore, Thailand, the Philippines, Indonesia and Malaysia) are the most likely candidates in Asia to foster financial cooperation and undertake measures to improve regional integration of financial markets.

Table 4.1: Stylized facts on Asian markets

	Mal.	Indo.	Thai.	Phil.	Sing.	HK	Jap.
Market cap. (2010)	408 689	360 388	277 731	157 320	647 226	2 711 316	3 827 774
KAOPEN (2011)	-1.168	-0.112	-1.168	-1.168	2.439	2.439	2.439

Notes: The domestic market capitalizations (in USD millions), extracted from the World Federation of Exchanges database, are those of the national stock market indexes included in our panel. *KAOPEN* is based on the binary dummy variables that codify the tabulation of restrictions on cross-border financial transactions reported in the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions*. According to Chinn and Ito (2008), *KAOPEN* This index takes on higher values the more open the country is to cross-border capital transactions.

These markets are likely to response in different ways to shocks depending on whether they are integrated with world or regional markets, that is the US or Japan (see Ng , 2000). To investigate the relative importance of the world's two largest stock markets on smaller Asian stock markets, we include the Standard & Poor's 500 (SPX) to our panel. Graphics 4.1 and 4.2 reproduce the respective time paths of range-based volatility on the emerging and developed stock markets. Clearly, the volatilities represented on Graphic 4.2 seem to be highly related when comparing with those of the Graphic 4.1.[8]

4.2.2 Econometric framework

Considering the stylized fact that shocks in volatility persist over time (see, e.g., Berger et al. , 2009, for a discussion on the source of the volatility persistence), the cointegration techniques are useful to guard against the risk of spurious regression and investigate whether volatility processes share a

6. Data are obtained from Bloomberg.
7. The MSCI and FTSE group classify Singapore and Hong Kong as developed markets.
8. The turmoil of December 2006 for SET is due to the currency reserve requirement imposed by the government on foreign stock investments.

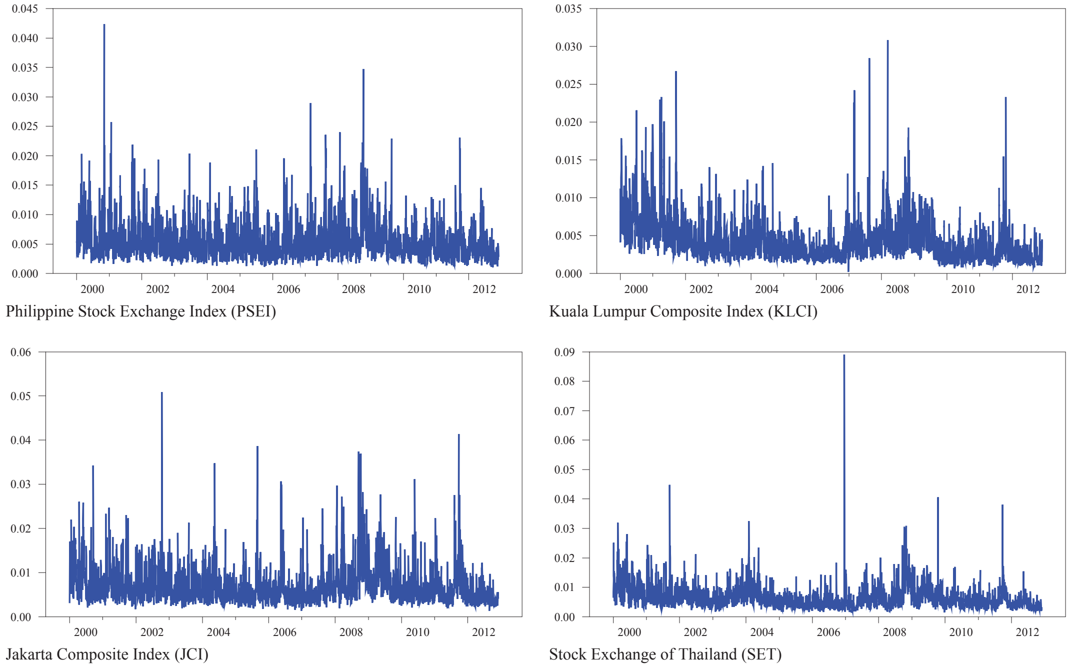

Figure 4.1: Daily range of emerging stock market indexes.

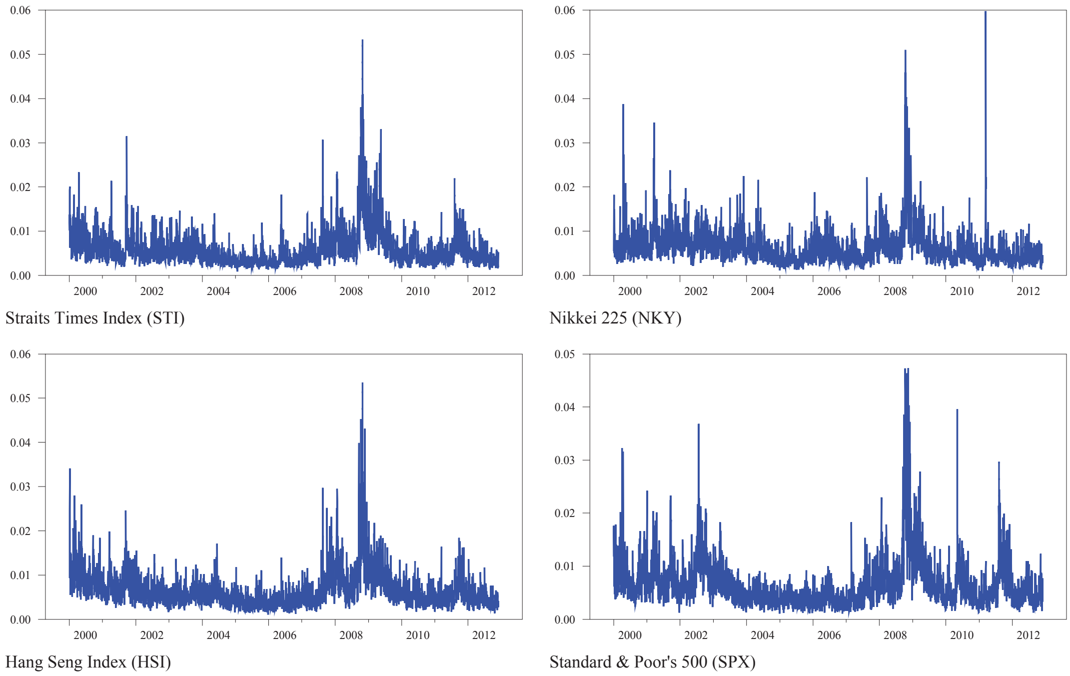

Figure 4.2: Daily range of developed stock market indexes.

common stochastic trend.[9] Indeed, cointegration is a powerful theory devoted to the analysis of long-run relationship. It states that a vector of p series, X_t, integrated of same orders, share $p - r$ common stochastic trends if there exists r linear combinations between the elements of X_t, having smaller orders of integration than X_t. The concept initiated by Granger (1981) does not restrict integration orders of X_t and the cointegrating errors to be integers. Accordingly, cointegration with real integration orders, termed fractional cointegration, has attracted a growing attention during the last decade and notably the interesting case where X_t is covariance-stationary but long-range dependent. Among others, recent contributions of Nielsen (2007), Robinson (2008a) and Shimotsu (2012) can be listed concerning univariate case while Hualde and Robinson (2010) and Johansen and Nielsen (2012) treat the multivariate case. This concept of stationary fractional cointegration is particularly relevant in empirical finance in which volatility series are often found to be covariance-stationary but exhibit long memory (see, e.g., Andersen et al. , 2003; Christensen and Nielsen , 2006). Nonetheless, there is no theoretical reason for the volatility process to be stationary. Using more appropriate techniques, Berger et al. (2009) and Frederiksen et al. (2012) produce estimates of the volatility persistence in the non-stationary region. Therefore, we are interested by estimators of fractional cointegration able to assess both stationary and moderately non-stationary processes.

4.2.3 Co-persistence model

We consider a simple cointegration system to investigate pairwise long-run relationships between volatility of market a and b,

$$(1-L)^{\gamma}(r_{t,a} - \beta r_{t,b}) = \varepsilon_{1t}, \qquad (1-L)^{\delta} r_{t,b} = \varepsilon_{2t}, \qquad t = 1, 2, ..., n, \qquad a \neq b, \tag{4.1}$$

In Eq. (4.1), $\{r_{t,.}\}$ is the low frequency daily range defined by the difference between the highest $(h_{t,.})$ and the lowest $(l_{t,.})$ log security prices over a day[10]. Terms $(1-L)^{\delta}$ and $(1-L)^{\gamma}$ are the fractional filter, further denoted Δ^{δ} and Δ^{γ} respectively. Parameters δ captures the persistence nature of the two observed volatilities while parameter γ measures the persistence of the long-run residuals ε_{1t}. According to the cointegration theory, co-persistence occurs when the long-run coefficient β is non-null and the strength of the long-run relation, $\nu = \delta - \gamma$, is positive.

4.2.4 The strategy of estimation

Fractional cointegration can generally be investigated following two different methodologies. The first is the so-called regression-based approach and requires to estimate the cointegrating vector(s), the integration orders of the regressor(s) and the cointegrating residuals. It has the advantage of providing accurate information concerning the strength of the long-run relationship. Nonetheless, it implies some difficulties in testing for cointegration when integration orders are not confined in a particular region of the parameter space. In such cases, there is little consensus on how to test for fractional cointegration. The second consists of estimating the cointegrating rank and does not imply any such requirement. However, it often requires tuning parameters that increases the sensitivity of the results and one can encounter some difficulties to obtain entirely conclusive results. To overcome these difficulties and to

9. Indeed, this risk exist whether or not the time series are covariance stationary as long as their integration orders sum up to a value greater than 1/2 (see Tsay and Chung , 2000).

10. Their is an extensive literature concerning realized volatility measures. In line with many studies, we support that the range-based volatility is a good proxy for the *true* volatility (see, e.g., Martens and Van Dijk , 2007; Engle et al. , 2012).

be more confident in the interpretation of the results, we adopt the following strategy: first, we test for the homogeneity of the integration orders δ_a and δ_b using an extended version of the Robinson and Yajima (2002)'s method that accommodates both stationarity and non-stationarity; second, we save the pairs for which $\delta_a = \delta_b$ and estimate their cointegrating rank by applying the rank estimator of Nielsen and Shimotsu (2007); third, we implement the regression-based approach by applying the exact local Whittle estimator of fractional cointegration proposed by Shimotsu (2012). [11]

All the aforementioned procedures operate in frequency domain and are semi-parametric thus focusing on a degenerating part of the periodogram around the origin. Accordingly, our approach is invariant to short-run dynamics and we are not concerned by misspecification issues. The later advantage has its counterpart since in consideration of the sensitivity of these procedure to the choice of the bandwidth, m, we have to be careful in interpreting the results. In this study, we are essentially concerned by the fact that bandwidth parameter must not tend very fast to ∞ with n to avoid higher frequency bias. This consideration is of importance because volatility is often subject to a perturbation term (see Frederiksen et al. , 2012), not modeled in our framework, which state the bandwidth requirement to $m = o(n^{2\delta/(1+2\delta)})$. In our case, we support that $m = \lfloor n^{0.5} \rfloor$ is reasonable considering both, this bandwidth requirement and the sample size $n = 3315$. Therefore, results are only reported for $m = \lfloor n^{0.5} \rfloor$ although several values for tuning parameters and bandwidths have been considered along the paper. [12]

4.2.5 Testing for equality of integration orders

Testing for the equality of δ_a and δ_b is a necessary condition for $r_{t,a}$ and $r_{t,b}$ to be non-trivially cointegrated. Considering the null hypothesis of pairwise equality, $H_0 : \delta_a = \delta_*, a = 1, \ldots, p, p = 2$, Robinson and Yajima (2002) propose the following test statistic

$$\hat{T}_0 = m(S\hat{\delta}_{ab})' \left(S\frac{1}{4}\hat{D}^{-1}(\hat{G} \odot \hat{G})\hat{D}^{-1}S' + h(n)^2 I_{p-1} \right)^{-1} (S\hat{\delta}_{ab}), \qquad p = 2, \tag{4.2}$$

where $\hat{\delta}_{ab} = (\hat{\delta}_a, \hat{\delta}_b)'$, I_p is an identity matrix of dimension p, $\odot$ denotes the Hadamard product, $D = \operatorname{diag}\left\{G_{11}, \ldots, G_{pp}\right\}$ and $S = [I_{p-1}; -\iota]$, with ι a $(p-1)$-vector of ones. $\hat{G}$ is the cross-spectral density matrix of residuals at zero-frequency. The tuning parameter $h(n)$ is chosen such as $h(n) = (\log n)^{-1}$, following Nielsen and Shimotsu (2007). Under the null hypothesis, $\hat{T}_0 \xrightarrow{d} \chi^2_{p-1}$ if r_t is not cointegrated and $\hat{T}_0 \xrightarrow{p} 0$ if r_t is cointegrated, with $r_t = (r_{t,a}, r_{t,b})'$.

Parameters $\hat{\delta}_a$, $\hat{\delta}_b$ and $\hat{G}$ are obtained from a multivariate version of the two-step exact local Whittle estimator (2S-ELW) of Shimotsu (2010) [13]. Table 4.2 reports the 2S-ELW estimates of persistence of each volatility for the bandwidth $m = \lfloor n^k \rfloor$, with $\lfloor . \rfloor$ the floor function and $k = \{0.4, 0.5, 0.6, 0.7, 0.8\}$. We observe that all long memory parameters are nor confined in a close interval neither regrouped in a precise region of the parameter space. Indeed, most of the volatilities are covariance stationary ($\delta < 0.5$) albeit in some cases series are moderately non-stationary ($\delta \geq 0.5$), encouraging us to consider an empirical strategy that accommodates both stationary and non-stationary cases.

11. Nielsen and Frederiksen (2011) and de Truchis and Keddad (2013) employ a similar methodology applied to stationary and non-stationary data respectively.

12. All unreported results are available upon request.

13. Given that the 2S-ELW is consistent for $\delta \in (-1/2, 2)$ and has a normal limit distribution $\sqrt{m}(\hat{\delta} - \delta) \xrightarrow{d} N(0, 1/4)$, it accommodates both the stationary and the non-stationary regions of the parameter space.

Table 4.2: 2S-ELW estimates of δ with $m = \lfloor n^k \rfloor$

k	SET	STI	PSEI	KLCI	JCI	HCI	NKY	SPX
0,4	0,457	0,473	0,257	0,682	0,180	0,584	0,478	0,605
0,5	0,373	0,531	0,248	0,402	0,311	0,676	0,425	0,586
0,6	0,451	0,602	0,262	0,399	0,366	0,657	0,597	0,605
0,7	0,363	0,557	0,281	0,363	0,340	0,535	0,474	0,620
0,8	0,362	0,455	0,304	0,397	0,387	0,430	0,433	0,509

4.2.6 The rank analysis

The test of Robinson and Yajima (2002) is crucial in view of determining the existence of a common value for δ_a and δ_b although this value remains unknown. Following Nielsen and Shimotsu (2007), we trivially estimate this common value, termed δ_*, by $\bar{\delta}_* = p^{-1}\sum_{a=1}^{p}\hat{\delta}_a, p = 2$ (see Table 4.3).[14] Then, we estimate $G(\delta_*)$ by substituting δ_* with $\bar{\delta}_*$. The bandwidth m_1 is introduced in estimating the matrix $G(\delta_*)$ to ensure that $\bar{\delta}_*$ converges to δ_* at a faster rate than $\hat{G}(\bar{\delta}_*)$ to $G(\bar{\delta}_*)$. We can employ both, the matrix $\hat{G}(\bar{\delta}_*)$ and the correlation matrix $\hat{P}(\bar{\delta}_*) = \hat{D}(\bar{\delta}_*)^{-1/2}\hat{G}(\bar{\delta}_*)\hat{D}(\bar{\delta}_*)^{-1/2}$ with $\hat{D}(\bar{\delta}_*) = \operatorname{diag}\left\{\hat{G}_{11}(\bar{\delta}_*), \ldots, \hat{G}_{pp}(\bar{\delta}_*)\right\}, p = 2$, to estimate the rank. Denoting τ_i the i-th eigenvalue of $\hat{G}(\bar{\delta}_*)$ or $\hat{P}(\bar{\delta}_*)$, the rank estimate is defined as

$$\hat{r} = \arg\min_{u=0,\ldots,p-1} L(u), \qquad L(u) = \upsilon(n)(p-u) - \sum_{a=1}^{p-u}\tau_a, \qquad p = 2 \tag{4.3}$$

Equation of $L(u)$ embed an additional tuning parameters, $\upsilon(n)$, for which the procedure is highly sensitive. In our application, we employ $\upsilon(n) = m_1^{-k}$ with $k = \{0.35, 0.25\}$. Simulations of Nielsen and Shimotsu (2007) reveal that for high values of k, the procedure is more likely to provide underestimate of $\hat{r}$ while for small values of k, the procedure is more likely to provide overestimate of $\hat{r}$.

Table 4.3: Empirical estimates of δ_* based on 2S-ELW estimates with $m = \lfloor n^{0.5} \rfloor$

$\bar{\delta}_*$	STI	PSEI	KLCI	JCI	HSI	NKY	SPX
SET	0,452	0,311	0,387	0,342	0,524	0,399	0,479
STI		0,390	0,466	0,421	0,603	0,478	0,559
PSEI			0,325	0,280	0,462	0,337	0,417
KLCI				0,357	0,539	0,413	0,494
JCI					0,494	0,368	0,449
HSI						0,550	0,631
NKY							0,505

4.2.7 The regression-based approach

The aforementioned rank analysis procedure does not provide information about the strength of the long-run relationship. Accordingly, the regression-based approach is particularly interesting to go further in the cointegration analysis. In view of anticipating the possible non-stationarity of the observed

14. We cannot perform a constraint estimation of the 2S-ELW because $G(\delta_a = \delta_b = \delta_*) = \frac{1}{m_1}\sum_{j=1}^{m_1} I_{\Delta(L;\delta_*,\delta_*)r}(\lambda_j)$ does not have full rank in presence of cointegration, with $\Delta(L;\delta_*,\delta_*)r = (\Delta^{\delta_*}r_{t,a}, \Delta^{\delta_*}r_{t,b})'$ and $\lambda_j = 2\pi jn^{-1}$ are the Fourier frequencies.

series (i.e. $\delta \geq 1/2$) we apply the exact local Whittle estimator of fractional cointegration (ELW-FC) developed by Shimotsu (2012). This estimator jointly estimates all the parameters of interest, β, γ and δ (henceforth denoted θ). It operates in two-steps to improve efficiency.

The first stage consists of applying a tapered version of the local Whittle (LW) estimator of Robinson (2008a) to estimate the cointegrated systems.[15] Let θ^d regroups long memory parameters such as $\theta^d = (\gamma, \delta)'$. Then, the objective function of the tapered LW estimator is defined by,

$$R(\theta) = \log\det\hat{\Omega}(\theta) - 2(\gamma+\delta)\frac{q}{(1-\kappa)m}\sum_{j(q,\kappa)}^{m}\log\lambda_j, \tag{4.4}$$

$$\hat{\Omega}(\theta)\frac{q}{(1-\kappa)m} = \sum_{j(q,\kappa)}^{m}\mathrm{Re}\left[\Psi(\lambda_j;\theta^d)BI_r(\lambda_j)B'\Psi(\lambda_j;\theta^d)^*\right], \qquad B = \begin{pmatrix} 1 & -\beta \\ 0 & 1 \end{pmatrix} \tag{4.5}$$

where $\Psi(\lambda_j;\theta^d) = \mathrm{diag}\left(\lambda^{\gamma}, \lambda^{\delta}e^{-i(\pi-\lambda_j)(\delta-\gamma)/2}\right)$, $I_r(\lambda_j)$ is the tapered periodogram matrix of r_t, $\lambda_j = 2\pi jn^{-1}$ are the Fourier frequencies and $\sum_{j(q,\kappa)}^{m}$ is the sum taken over $j = q, 2q, \ldots, m$ for $j \leq \lfloor\kappa m\rfloor$ and q the order of the taper. Conversely to the original LW estimator, the phase parameter is not estimated but modeled as $(\delta-\gamma)\pi/2$. Shimotsu (2012) also introduces a trimming parameter, κ, to control the behavior of the objective function when $\delta - \gamma > 1/2$. In the context of our application this trimming parameter take the following values $\kappa = \{0.05, 0.04\}$.

The second stage consists of an application of the 2S-ELW of Shimotsu (2010) to obtain θ^* as $\theta^* = \arg\min_{\theta\in\Theta} R^*(\theta)$, given the concentrated objective function,

$$R^*(\theta) = \log\det\tilde{\Omega}^*(\theta) - 2(\gamma+\delta)\frac{1}{m}\sum_{j=1}^{m}\log\lambda_j, \tag{4.6}$$

$$\tilde{\Omega}^*(\theta) = \frac{1}{m}\sum_{j=1}^{m}\mathrm{Re}\left[I_{\Delta^{\theta^d}r}(\lambda_j;\beta)\right], \qquad I_{\Delta^{\theta^d}r}(\lambda_j;\beta) = \omega_{\Delta^{\theta^d}r}(\lambda_j;\beta)\bar{\omega}_{\Delta^{\theta^d}r}(\lambda_j;\beta) \tag{4.7}$$

$$\omega_{\Delta^{\theta^d}r}(\lambda_j;\beta) = \begin{pmatrix} \omega_{\Delta^{\gamma}(r_{t,a}-\beta r_{t,b})}(\lambda_j) \\ \omega_{\Delta^{\delta}r_{t,b}}(\lambda_j) \end{pmatrix} \tag{4.8}$$

with Θ the parameter space of θ^* and $\omega_r(\lambda_j)$ the discrete Fourier transform of r_t.

Denoting $\hat{\theta}$ the estimate of θ from aforementioned tapered LW estimator, the 2S-ELW-FC estimator of Shimotsu (2012) is defined as $\theta^* = \hat{\theta} - \left((\partial^2/\partial\theta\partial\theta')R^*(\hat{\theta})\right)^{-1}\left((\partial/\partial\theta)R^*(\hat{\theta})\right)$.[16] The author shows that $m^{1/2}\mathrm{diag}\left\{\lambda_m^{-v_0}, 1, 1\right\}(\theta^* - \theta_0) \xrightarrow{d} N(0, \Xi^{-1})$ as $n \to \infty$ when $v_0 \in (0, 1/2)$ and $m^{1/2}(\theta^{d*} - \theta_0^d) \xrightarrow{d} N(0, \Xi_{\theta^d}^{-1})$ while $(\beta^* - \beta_0) = O_p(n^{-v_0})$ as $n \to \infty$ when $v_0 \in (1/2, 3/2)$ with v_0 the true value of $v = (\delta - \gamma)$. Exploiting $\tilde{\Omega}^*(\theta)$, this procedure is also able to estimate the off-diagonal parameter, ρ, of the residuals covariance matrix (i.e. endogeneity parameter).

15. See Velasco (1999) for a discussion about tapered estimators.

16. According to the author, an iterative procedure may result from estimation of θ^* without modifying the asymptotic distribution and may improve the finite sample properties.

4.3 Cointegration analysis

Table 4.4 reports the test statistic $\hat{T}_0$ for the 2S-ELW estimates of the pairwise fractional integration orders (see Table 4.2). If the null hypothesis is accepted at conventional significance levels, the pairwise integration orders are equals. Since the 95% critical value of the χ^2_1 distribution is 3,84 we easily accept, in most cases, the null of equality of the integration order. At this stage, it is worth noting that the null hypothesis is always accepted at a 99% level of significance excepted for nine pairs including emerging with some developed markets. This suggest nonidentical, distinct or independent volatility process between the emerging and developed markets. Regarding equality of integration orders among developed markets, the null hypothesis is always accepted with the only exception of Japan with Hong Kong, while the volatility processes of emerging markets always exhibit a common order of fractional integration. Accordingly, the fractional cointegration hypothesis can be investigated for 19 pairs of volatility series. [17]

Table 4.4: Test for the homogeneity of integration orders with $m = \lfloor n^{0.5} \rfloor$ and $h(n) = (\log n)^{-1}$

T_0	STI	PSEI	KLCI	JCI	HSI	NKY	SPX
SET	3,126	2,063	0,101	0,506	11,854	0,308	5,321
$\phi_\chi(T_0)$	(0,077)	(0,151)	(0,751)	(0,477)	(0,001)	(0,579)	(0,021)
STI		11,130	2,242	8,161	5,032	2,227	0,708
$\phi_\chi(T_0)$		(0,001)	(0,134)	(0,004)	(0,025)	(0,136)	(0,400)
PSEI			2,941	0,519	21,675	3,621	13,406
$\phi_\chi(T_0)$			(0,086)	(0,471)	(0,000)	(0,057)	(0,000)
KLCI				1,057	8,832	0,060	3,798
$\phi_\chi(T_0)$				(0,304)	(0,003)	(0,806)	(0,051)
JCI					19,249	1,687	10,048
$\phi_\chi(T_0)$					(0,000)	(0,194)	(0,002)
HSI						9,660	1,520
$\phi_\chi(T_0)$						(0,002)	(0,218)
NKY							4,613
$\phi_\chi(T_0)$							(0,032)

Notes: P-values are displayed in parentheses (.).

Following Nielsen and Shimotsu (2007), the rank estimates are conducted by using the correlation matrix $\hat{P}(\bar{\delta}_*)$. The results are presented in Table 4.5. In the bivariate case u takes the value 0 or 1 depending on the rank estimate. We adopt the regression-based approach to confirm the rank analysis and give more intuitive results about the strength of the long-term relationships. Table 4.6 reports the results of the ELW-FC estimates for $m = \lfloor n^{0.5} \rfloor$ and $\kappa = 0.05$. [18]

According to the rank analysis, the results indicate the presence of several cointegration relationships. Indeed, for the case with the lowest $\upsilon(n)$, i.e. $\upsilon(n) = m_1^{-0,25}$, we find eleven pairs for which the volatility processes are co-persistent. Again, we find a strong heterogeneity among the sample. For instance, the volatility of the emerging stock markets (i.e. PSEI, KLCI, SET, and JCI) appear to be well tied together while the share of commonality with their regional (STI, HSI and JPY) and global (SPX) counterparts is less evident. Conversely, HSI, STI are pairwise fractionally cointegrated both with each

17. We choose the regressor according to the market capitalization (see Table 4.1).

18. Our results are not sensitive to the trimming parameter κ. However, we apply a penalty parameter to the coefficient $\hat{\beta}$ for the pairs HSI-SPX and KLCI-NKY because of convergence considerations of the concentrated objective function (see Shimotsu , 2012).

Table 4.5: Pairwise rank estimates based on $\hat{P}(\bar{\delta}_*)$

	SET/SPX		KLCI/SPX		STI/SPX		HSI/SPX		NKY/SPX	
$\upsilon(n)$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$
$L(0)$	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194
$L(1)$	-1,240	-1,117	-1,196	-1,074	-1,590	-1,467	-1,523	-1,400	-1,497	-1,374
$\hat{r}$	0	0	0	0	1	1	1	1	1	1
	PSEI/NKY		SET/NKY		JCI/NKY		KLCI/NKY		STI/NKY	
$\upsilon(n)$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$
$L(0)$	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194
$L(1)$	-1,237	-1,114	-1,215	-1,092	-1,335	-1,212	-1,224	-1,101	-1,538	-1,415
$\hat{r}$	0	0	0	0	0	1	0	0	1	1
	STI/HSI		SET/STI		KLCI/STI		PSEI/KLCI		SET/KLCI	
$\upsilon(n)$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$
$L(0)$	-1,440	-1,194	-1,440	-1,194	-(1,440)	-1,194	-1,440	-1,194	-1,440	-1,194
$L(1)$	-1,599	-1,476	-1,297	-1,174	-1,351	-1,228	-1,298	-1,175	-1,255	-1,132
$\hat{r}$	1	1	0	0	0	1	0	0	0	0
	JCI/KLCI		PSEI/JCI		SET/JCI		PSEI/SET			
$\upsilon(n)$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$	$m_1^{-0,35}$	$m_1^{-0,25}$		
$L(0)$	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194	-1,440	-1,194		
$L(1)$	-1,325	-1,202	-1,331	-1,208	-1,353	-1,230	-1,347	-1,224		
$\hat{r}$	0	1	0	1	0	1	0	1		

other and SPX, while NKY is pairwise fractionally cointegrated with only STI and SPX. As mentioned above, the procedure is sensitive to the tuning parameter $\upsilon(n)$, and co-persistence among emerging markets disappear when $\upsilon(n) = m_1^{-0,35}$. Overall, it follows that only developed stock markets, i.e. HSI, STI, NKY, and SPX, are pairwise fractionally cointegrated (except for the pair including HSI and NKY since their integration orders are not equals) and this finding is robust to the use of different tuning parameter values.

The estimates of β which are not significants at conventional levels imply the absence of long-run co-movement between the two markets. Consequently, KLCI is not fractionally cointegrated with NKY, STI, PSEI, SET, and JCI while, SET, PSEI and JCI are not pairwise fractionally cointegrated with each other. Taking as a whole, these results are consistent with the conclusions drawn on the basis of the regression analysis., except for the pair including STI and NKY for which the estimate of β is insignificant. In such a case, we must be careful when interpreting the estimation results.

Interestingly, the estimates of β are larger among developed markets in most cases, suggesting again a strong correlation among their volatility processes. For instance, correlations between SPX and STI, HSI and NKY, are equal to 0.769, 0.641 and 0.723, respectively. We find also a large $\hat{\beta}$ for STI-HSI (0.91), SET-JCI(1.309) and SET-KLCI(1.160). Conversely, the long-term correlations among the emerging and developed markets are less than one-half in most cases.

Another necessary condition to establish a fractional cointegration relationship between two series is $\hat{\gamma} < \hat{\delta}$. In most cases, the cointegrating residuals have substantially smaller memory parameters estimates than $\hat{\delta}$, suggesting fractional cointegration between these pairwise volatility series. In addition, the gap between the estimates of $\hat{\delta}$ and $\hat{\gamma}$, provides informations about the persistence of deviations from the long-run equilibrium. The higher the gap between δ and γ, the greater the strength of the

Table 4.6: ELW-FC estimates for $m = \lfloor n^{0.5} \rfloor$ and $\kappa = 0.05$

	SET/SPX		KLCI/SPX		STI/SPX		HSI/SPX		NKY/SPX	
$\hat{\beta}$	0,429	(0,090)	0,341	(0,092)	0,769	(0,128)	0,641	(0,107)	0,723	(0,105)
$\hat{\gamma}$	0,295	(0,066)	0,386	(0,065)	0,426	(0,066)	0,501	(0,065)	0,349	(0,066)
$\hat{\delta}$	0,602	(0,066)	0,606	(0,065)	0,588	(0,066)	0,700	(0,065)	0,584	(0,066)
$\hat{\rho}$	-0,150		-0,254		-0,187		-0,282		-0,158	
	PSEI/NKY		SET/NKY		JCI/NKY		KLCI/NKY		STI/NKY	
$\hat{\beta}$	0,160	(0,068)	0,575	(0,228)	0,431	(0,109)	0,370	(0,446)	0,794	(0,741)
$\hat{\gamma}$	0,131	(0,066)	0,256	(0,065)	0,126	(0,066)	0,381	(0,066)	0,465	(0,066)
$\hat{\delta}$	0,427	(0,066)	0,422	(0,065)	0,430	(0,066)	0,439	(0,066)	0,425	(0,066)
$\hat{\rho}$	0,183		-0,214		0,132		-0,171		-0,087	
	STI/HSI		SET/STI		KLCI/STI		PSEI/KLCI		SET/KLCI	
$\hat{\beta}$	0,910	(0,035)	0,605	(0,120)	0,301	(0,188)	0,428	(0,112)	1,160	(0,316)
$\hat{\gamma}$	0,244	(0,066)	0,265	(0,066)	0,401	(0,066)	0,108	(0,066)	0,210	(0,065)
$\hat{\delta}$	0,678	(0,066)	0,537	(0,066)	0,523	(0,066)	0,401	(0,066)	0,402	(0,065)
$\hat{\rho}$	0,101		-0,120		0,194		0,036		-0,290	
	JCI/KLCI		PSEI/JCI		SET/JCI		PSEI/SET			
$\hat{\beta}$	0,616	(0,387)	0,277	(0,187)	1,309	(0,451)	0,251	(0,110)		
$\hat{\gamma}$	0,237	(0,065)	0,194	(0,066)	0,213	(0,057)	0,170	(0,065)		
$\hat{\delta}$	0,403	(0,065)	0,310	(0,066)	0,305	(0,057)	0,371	(0,065)		
$\hat{\rho}$	0,232		0,154		-0,719		0,262			

Notes: Standard deviations are displayed in parentheses (.).

cointegrating relation. The values $\hat{\delta} - \hat{\gamma}$ are particularly high for the cointegrated pair including Hong Kong with Singapore (0.435). We also observed some interesting relationships between emerging and developed markets, such as PSEI-NKY (0.296), JCI-NKY (0.304), SET-SPX(0.308). Again, we need to be careful about the relevancy of these findings since this procedure is sensitive to the bandwidth selection and other tunning parameters.[19] Nonetheless, these last findings seem to confirm that the developped markets, especially Hong Kong and Singapore, are closely linked in the long run with both each other and the US stock market.

4.4 Concluding remarks

In this study, we examine the integration of stock markets in several emerging and developed Asian markets. Conversely to most of the existing studies that investigate the short-term correlation, we focus on long-term cross-border correlation of integrated volatilities. More precisely, we investigate the persistent and the co-persistence nature of volatility processes by means of fractional cointegration techniques. We implement a robust methodology that relies on both, the rank estimate and the regression-based analysis. Our findings confirm that developed stock markets in Asia are globally integrated with the world stock market of the US, consistently with the view expressed by Dao and Wolters (2008)

19. For robustness, we also study a parametric form of the model in Eq. (4.1) using the quasi-maximum likelihood (QML) estimator of de Truchis (2013). According to this investigation, short run dynamics of the cointegrating errors are fairly complex and encourage us to consider small bandwidths. We find that results are not so different as regards the relation of interest, although the QML estimator seems to widely underestimate β.

who concluded that volatilities in Singapore, Hong Kong, Japan an US stock markets are in essence co-persistent. Accordingly, volatilities in the US and developed markets in Asia respond in a similar way to shocks, implying greater co-movements in the long run among their volatility process. This also implies that effective diversification of portfolios among these markets cannot be achieved in the long run. It is worth noting that empirical results concerning the Japanese stock market are mixed, since only the pair including the SPX is robust to the two fractional cointegration methods. This means that the Japanese stock market is integrated with the world stock market of the US but not with those of Hong Kong and Singapore. The fact that Japan holds a sizable share of its total foreign assets in the form of US assets (32.3 % in 2009) and invests very little in Asia (2.4% in 2009) may provide evidence of such an integration with the US stock market (see Park and lee , 2011 for details on the pattern of cross-border financial asset holdings in Asia).

Conversely, there is only little evidence of co-persistence among the emerging markets (i.e. PSEI, JCI, KLCI and SET), which appear to be segmented at both regional and global levels over the 2000-2012 period. This suggests that these markets do not share in the long run commonalities in the market fundamentals underlying their volatility processes. These results are consistent with Harvey (1995) and Beckaert and Harvey (1995), who argue that emerging markets are more likely to be influenced by local events. Yu et al. (2010) explain the divergence in integration by the lack of harmonization of standards in capital markets and the absence of links between jurisdictions across the whole spectrum of financial infrastructure. Therefore, the emerging markets in Asia need to address these issues in the planning of a coordinated strategy for promoting financial integration within the region, along with their political support toward greater financial cooperation. This would help the region to improve resilience against country-specific shocks via the cross-ownership of equity capital. On the other hand, from the point of view of individual investors who seek to diversify in these emerging markets, the little evidence of co-persistence indicates that, in the long run, the potential benefits of international portfolio diversification remain effective.

Conclusions et perspectives

Nous souhaitions, dans cette thèse, évaluer le degré d'intégration monétaire et financière en Asie. Étant donné que les pays de l'ASEAN-5 se présentent comme les candidats les plus sérieux pour initier une nouvelle étape dans la mise en oeuvre d'une intégration monétaire plus aboutie, notre recherche s'est principalement concentrée sur l'étude de ces cinq pays.

Depuis la crise de 1997-98, les autorités politiques et économiques s'interrogent sur la nécessité d'adopter une nouvelle stratégie de change collective afin de parvenir à une plus grande stabilité des taux de change intra-régionaux. La volonté de coopérer plus étroitement dans ce domaine s'est manifesté en 2006, à l'occasion de la 39ème réunion annuelle de la Banque Asiatique de Développement, par le désir de créer une unité monétaire asiatique (*Asian Currency Unit*, ACU). Dans un premier temps, son objectif est exclusivement orienté vers la surveillance du mouvement relatif des devises asiatiques mais, à terme, l'ACU pourrait jouer un rôle plus important dans la coordination des politiques de change. A cet égard, nous nous sommes intéressé à l'importance de l'ACU dans la gestion des politiques de change en Asie. Nous avons montré que le dollar continuait de prédominer dans le panier de devises de ces pays, suggérant un rôle limité de l'ACU avant 2006. Néanmoins, après 2006, l'influence du dollar semble décroitre au profit de l'ACU et de l'euro. Ce résultat met en évidence leur volonté de se détourner d'une politique de change exclusivement centrée sur le dollar, vers une politique plus flexible, avec un panier de devises comme point d'ancrage. Néanmoins, il serait faux d'affirmer que les pays asiatiques ont manifesté ces dernières années une volonté explicite de coordonner leur politique de change. Au contraire, le poids persistant du dollar peut traduire un défaut de coordination dans le choix du régime de change optimal au sein de la région. En effet, aussi longtemps que les pays asiatiques continuent d'ancrer leur monnaie sur le dollar, aucun pays asiatique n'aurait intérêt à se détourner unilatéralement de celui-ci s'il souhaite préserver sa compétitivité et promouvoir ses exportations. Dans cette optique, nous avons montré que la décision des autorités monétaires chinoises de se détourner d'un ancrage exclusif sur le dollar, au profit d'un panier de devise, pouvait avoir encouragé les pays asiatiques à introduire davantage de flexibilité dans la gestion de leur taux de change.

Si la coordination des politiques de change peut constituer une étape préalable à l'adoption d'une monnaie unique, la constitution d'une union monétaire suppose néanmoins certains pré-requis. Depuis Mundell (1961), l'examen des conditions d'optimalité d'une zone monétaire en Asie a engendré une littérature volumineuse. Le débat a essentiellement porté sur la nature des chocs macroéconomiques et le degré de synchronisation des cycles économiques. Pour autant, aucun consensus n'a été atteint. C'est pourquoi, dans un deuxième chapitre, nous avons souhaité apporter un éclairage nouveau sur cette question en tenant compte de la non-linéarité dans la dynamique du cycle. Dans cette optique, nous avons utilisé un modèle Markov-switching dont les probabilités de transition dépendent du temps (Filardo , 1994). L'objectif de ce chapitre était double. D'une part, nous avons cherché à établir si les corrélations entre les cycles des affaires dépendent du régime du cycle (phases de contraction et de récession). D'autre part, l'objectif était de déterminer si ces corrélations provenaient de la transmission des cycles à l'échelle régionale (entre les pays de l'ASEAN-5) où si elles étaient le fruit de la dynamique des cycles globaux (Chine, Japon, USA) supposés jouer un rôle important dans la synchronisation au sein de la région. Nous avons utilisé un modèle Markov-switching dont les probabilités de transition varient dans le temps (Filardo , 1994) afin de modéliser la non-linéarité et la dynamique des corrélations. Nous avons montré que la corrélation entre les cycles était plus forte durant les phases de contraction mais que la dynamique d'ajustement restait propre à chaque pays. Par ailleurs, contrairement aux cycles globaux, certains cycles des affaires asiatiques contiennent des informations pertinentes pour prédire les changements de régime des autres pays, ce qui tendrait à prouver que l'ASEAN-5 est découplé du reste du monde. Néanmoins, nos résultats ont montré qu'une union monétaire parmi l'ASEAN-5 restait prématurée.

Afin d'étendre notre analyse sur la faisabilité d'une union monétaire en Asie, le troisième chapitre s'est concentré sur la cointégration fractionnaire des taux de change réels parmi l'ASEAN-5. Nous avons testé empiriquement la théorie de la parité des pouvoirs d'achat généralisée introduite par Enders and Hurns (1994, 1997). Selon les auteurs, lorsque les fondamentaux macroéconomiques influant sur l'évolution des taux de change réels sont similaires, il est possible que ces derniers partagent une tendance commune de long-terme. Néanmoins, les études existantes se sont uniquement focalisées sur la cointégration $I(1)/I(0)$ des taux de change réels. Considérant les limites de l'approche traditionnelle, nous avons supposé que les taux de change réels pouvaient être caractérisés par un processus fractionnairement cointégré. Nous avons identifié plusieurs relations de long-terme dites “faibles”, en ce sens que les résidus du système de cointégration étaient très persistants mais *mean-reverting*. Si une union monétaire entre tous les pays de l'ASEAN-5 semble compromise, une intégration monétaire plus poussée entre différents sous-groupes demeure possible. Dans la perspective d'une future union monétaire, nous avons également montré qu'un ancrage régional sur le dollar était moins désirable qu'un ancrage sur un panier de devises composé du yen et du dollar. Une possible extension serait d'étudier la cointégration fractionnaire des taux de change effectifs réels afin de s'affranchir du choix de la monnaie ancre.

Enfin, nous avons étudié le degré d'intégration financière en Asie au moyen de la cointégration fractionnaire dans la volatilité des marchés boursiers asiatiques. L'objectif était de montrer si les marchés boursiers en Asie étaient intégrés entre eux et/ou avec les marchés boursiers internationaux où si, au contraire, ils étaient segmentés les uns par rapport aux autres. Nous avons montré que les volatilités observées sur les marchés boursiers internationaux partageaient une tendance stochastique commune. Ce résultat est cohérent avec l'idée selon laquelle la volatilité des marchés boursiers internationaux est guidée par des facteurs globaux d'intégration, alors que les marchés boursiers des pays émergents apparaissent segmentés, tant au niveau régional qu'au niveau global. De ce point de vue, les marchés financiers des pays émergents asiatiques restent peu intégrés entre eux et avec le reste du monde, limitant ainsi la possibilité pour ces pays de lisser leur consommation et partager le risque macroéconomique face aux chocs spécifiques.

Dans l'ensemble, les résultats obtenus dans cette thèse montrent que l'adoption d'une monnaie unique en Asie n'est pas envisageable actuellement, même si une coopération monétaire plus étroite peut être souhaitable. De ce point de vue, il est important de noter que l'intégration monétaire est avant tout un processus graduel, où la création de mécanismes coopératifs est un préalable à une coordination plus active des politiques monétaires. Ainsi, une approche progressive débutant par la mise en place de politiques à l'échelle régionale, dont le niveau d'exigence en termes de coopération serait moindre, permettrait aux pays asiatiques de gagner en expérience avant d'évoluer vers une forme d'intégration monétaire plus avancée. Cette coopération monétaire pourrait viser dans un premier temps à stabiliser les taux de change intra-régionaux afin de favoriser les échanges commerciaux, la croissance et la stabilité macroéconomique régionale. La création de l'ACU pourrait constituer un modeste pas dans cette direction. Cette monnaie synthétique pourrait jouer le rôle de point d'ancrage au sein d'un mécanisme de change asiatique ou celui d'une monnaie parallèle (voir Eichengreen , 2006). Une autre stratégie consisterait à opérer un changement collectif de leur régime de change actuel vers un régime de change intermédiaire tel que l'adoption d'un panier dont les devises et pondérations seraient communes ou adaptées à chaque pays. À mesure que l'interdépendance macroéconomique se renforcerait au sein de la région, l'adoption d'une monnaie unique se ferait plus désirable. Par ailleurs, si l'objectif pour les pays asiatiques reste avant tout de stabiliser les taux de change intra-régionaux afin de promouvoir la croissance de leurs exportations, il convient alors de noter que cet objectif peut être atteint sans l'adoption d'une monnaie commune. À cet égard, l'adoption d'un ancrage sur un panier de devises dont la pondération serait relativement identique au sein de la région pourrait se révéler suffisant.

Il existe cependant plusieurs sources de scepticismes à l'égard d'une coopération monétaire plus étroite. La plus importante est liée au poids politique et économique de la Chine et du Japon au sein de la région, ainsi qu'à la rivalité lancinante qui les oppose pour atteindre l'hégémonie en Asie. Si une union monétaire devait se créer, il est certain que le yen et le yuan ne pourraient rester en marge du processus d'intégration monétaire. Cela semble d'autant plus problématique lorsque l'on admet, d'une part, le fait que le Japon et la Chine ne paraissent pas disposés à abandonner leur régime de change actuel et, d'autre part, l'asymétrie dans la conduite de leur politique de change respective. Par ailleurs, si une union monétaire devait être institutionnalisée autour du yen ou du yuan, il parait peu probable que l'un ou l'autre de ces pays ne soit disposé à abandonner sa politique monétaire au profit de l'autre. De ce point de vue, les relations conflictuelles entre le Japon et la Chine pourraient obstruer une coopération plus étroite au sein de la région. L'expérience européenne, récente comme plus ancienne, semble démontrer que l'intégration économique et monétaire ne peut rester en marge du processus politique. Si l'intégration économique en Asie émane principalement des forces du marché, l'Asie manque néanmoins d'un véritable esprit communautaire qui se veut indispensable dans l'élaboration d'une coopération monétaire et économique plus aboutie. Au delà des critères purement économiques, il est donc nécessaire de considérer la dimension culturelle, historique et politique qui peut lier ces pays (Kwack ; 2004, 2005; Plummer , 2006; Volz , 2006; Chey , 2009).

Il serait tout de même précipité de disqualifier la perspective d'une coopération monétaire régionale plus élaborée. En effet, l'Asie est une région en pleine mutation où des progrès ont été déjà accomplis dans les domaines de la coopération monétaire et financière. Des efforts importants ont été entrepris pour améliorer la surveillance économique régionale (*the Economic Review and Policy Dialog* où ERPD, et le groupe régional de banques centrales EMEAP), promouvoir le développement des marchés obligataires en monnaie locale (l'initiative *Asian Bond Markets* ABF2 et le Fonds Obligataire Asie ABMI) et établir un mécanisme d'échange de devises sous forme de *swaps* (l'Initative de Chiang Mai Multilatéralisé en 2009 CMIM). Ainsi, les dirigeants des pays asiatiques semblent être conscients de la nécessité de coopérer plus étroitement, en parallèle d'une interdépendance économique qui ne cesse de croître. De ce point de vue, il est important pour les dirigeants de ces pays de considérer l'intégration monétaire dans une perspective plus large, afin de dépasser les obstacles qui font face au développement économique et de créer un environnement macroéconomique stable, essentiel pour leur croissance économique.

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