Financial Integration between China and Asia Pacific Trading Partners: Parities Evidence from the First- and Second-generation Panel Tests

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March 2012

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ABSTRACT

This paper presents a joint investigation of the international parity conditions between China and her 13 major trading partners in the Asia Pacific over globalization era. Several advanced tests of unit root for univariate and panel series are utilized in the analyses. Our findings reveal that first, RIP holds stronger than PPP among APEC-China. Second, both parities tend to hold better as one move to the recent years, attributed not only to the financial liberalization process among APEC economies, but also to the Chinese trade policy and the regional commitment for the ASEAN+3+2+1 cooperation. Third, China and APEC have improved the ability to absorb regional shocks as indicated by the shortened half-life reported over time, especially when the post-Asia crisis era is included.

Keywords: PPP, RIP, Panel Unit Root Tests, Mean Reversion, Half-life, Financial Integration

JEL Classification: C3, F36, G15

1. Introduction

Unlike her neighboring countries in the East Asia, China's economic reform programs are relatively recent, attributed to the closed-door policy and centrally-planned economic system during 1950s-1970s. However, the affluent human capital and economic resources has provided China the new impetus to and reinvigorate the economic reforms since 1978 and the economic progress of this economy is eye-catching. Within three decades, China has transformed itself from a rigid central-planning system to an increasingly open and market-oriented economy, with the achievement of averagely 9.7% real GDP growth per annum. As of November 2007, China recorded a nominal GDP of US$3.42 trillion and holds the fourth largest economy after the US, Japan and Germany. China's GDP officially overtook Japan in the second quarter of 2010 although the GDP of per capita ($8394) is still significantly lower than that of Japan (US$39731) and United Sates (US$46380).

China's role in the global trading and finance has steadily grown, especially after the accession to WTO in November, 2001. China is presently the world's largest exporter and second largest importer. In 2010, China's total trade exceeded US$2.8 trillion[1] and its current account surplus amounted to US$0.2 trillion, which ranked top globally (Data Stream). Despite being the major trading partners for many of the Asia Pacific economies (APEC)[2], China has also actively involved with the Chiang Mai Initiative (2000), the Bali Accord (2003) and the Singapore Declaration (2007) and devoted for closer cooperation within the ASEAN+3+2+1 framework. Additionally, China's efforts toward regionalism in most of the countries under review (in particular East Asian) that started in the last decades are expected to have some impact on her integration process with the APEC countries (see Yu, 2011). In line with the trade and exchange rate liberalization[3], China has gradually opened up the financial markets by permitting a wide variety of private enterprise in services and light manufacturing; developing a more diversified banking system and capitalized stock market; and increasing the foreign investments. According to the World Bank statistics, China has doubled her accumulated FDI since 1999 from US$39 billion, to around US$574 billion in 2010, to become the largest FDI destination in East Asia. Besides the European counterparts, Hong Kong, Taiwan, Japan and the US hold the major shares of foreign investments in China. Similarly, half of the stocks of foreign bank lending are also sourced from the US and East Asian trading partners.

Taking account of these developments mentioned above, markets convergence and future economic events that anchored by China are well expected in the Asia Pacific region. Yet, to what extent has China truly integrated with the regional economies, remains as major apprehension. Recent proposal of Trans-Pacific Partnership (TPP) negotiations on regional trade arrangements has elevated further debates among scholars (see Armstrong, 2011), since the 2011 Honolulu APEC meeting. Two unsolved but essential questions thus arise. First, is

[1] More than hundred times the total trade figure of US$20.6 billion in 1978.

[2] The directions of trade for selected APEC are accounted for 61% and 59% of the Chinese total exports and imports respectively in 2006. These selected APEC include the US, Australia, New Zealand, Taiwan, South Korea, Singapore, Indonesia, Malaysia, the Philippines and Thailand.

[3] China has repeatedly devalued its currency as a means of trade expansion and external competitiveness gains in the 1980s and the early 1990s. In 1994, 1996 and 2005, unification of multiple rates and liberalization of exchange rates drive the RMB a step further toward the full convertibility. Likewise, the portion of foreign trade under direct administrative controls has been substantially reduced while more subject to the market forces.

regional trade competition sufficient to eliminate prices arbitrage and hence reflecting the exchange value of Chinese Yuan when more and more trading of goods and services are promoted across borders? Second, are China's pricing and investment structures integrated with the regional standards to facilitate cross-border financial assets substitutability or allowed for greater portfolio diversification? The former question relates to the Purchasing Power Parity (PPP hereinafter) while the latter directs to the Real Interest Rate Parity (RIP hereinafter) condition.

Without the answers to the questions, we are unable to draw any conclusive conclusion about the extent of economic integration between China-APEC, and hence intricate the formulation of regional monetary and exchange rate policy coordination[4]. Yet, the empirical evidences of PPP and RIP, which have hitherto been abundant, are still contentious especially among developing economies (see Rogoff, 1996; Taylor and Taylor, 2004; Cheung, *et al.*, 2005; in recent surveys). Moreover, the assessment of parity conditions based on China-denominated exchange rates and financial securities are notably lacking and inconclusive. Among the few China-based studies, Finke and Rahn (2005) and Coudert and Couharde (2007) revealed that Chinese yuan significantly deviates from PPP, whereas Gregory and Shelley (2011) found evidence of PPP – only for the real effective yuan but not for the real yuan/USD rates. Cheung et al. (2003), in a separate endeavor, examined three parity conditions (PPP, UIP, RIP) consecutively and concluded that parities hold among China-Taiwan-Hong Kong. Chan et al (2012) then conducted a structural system to assess PPP and UIP for China-Japan. They confirm that both parity conditions hold in the long run when structural breaks of Asia crisis, subprime crisis and six over-identifying restrictions were taken into accounts. Meanwhile, Cavoli et al. (2004) examined the parity conditions for China, East Asia and ASEAN but failed to find clear indication of intensified financial integration. Likewise, Laurenceson (2003) shows that China-ASEAN's financial linkages remain weak though the market integration of goods and services is relatively well-established.

This paper aims to jointly investigate the validity of PPP and RIP conditions for China vis-à-vis her 13 trading partners in the Asia Pacific region. Such practice of joint investigation is not frequently applied in the literature but supported by Cheung et al. (2003) and Cavoli et al. (2004), among the few others. A different but clearer insight or perspective may be gained from the joint assessment of China and APEC emerging economies with different regulatory regimes at different stages of development. More important, monetary and exchange rate coordination policies derived from the PPP and RIP conditions within similar time zone would enable the Asia Pacific region to exert an important influence upon the future evolution of the global trade and financial system.

To assess PPP and RIP, a convenient strategy is to scrutinize the mean-reversion behaviors of bilateral real exchange rates (REX) and real interest differentials (RID) among China-APEC. Monthly observations and sub-samples within 1986-2007 are being considered to accentuate the effects of institutional changes and financial crisis, both local and regional. Due to the deficiency in extant econometric tests, various estimation methods are adopted to increase the

[4] Support for PPP would imply the goods market integration attributed to price convergence and apposite alignment of exchange rate, or otherwise. Similarly, acceptation of the RIP will uphold the regional financial integration among China-APEC while rejection of RIP may imply the greater degree of monetary autonomy.

likelihood of establishing well-defined results. These include the endogenous break test advocated by Saikkonen and Lütkepohl (2002), the first-generation panel tests by Levin-Lin-Chu (2002) and Im-Pesaran-Shin (2003) as well as the second-generation panel test by Pesaran (2007) that allow for cross-sectional dependency. Results of univariate and panel tests are compared in considering of the robustness within the macro-panel setting. To capture the degree of shock adjustments towards equilibrium, we also construct the half-life and confidence intervals by means of the correction factor model put forward by Rossi (2005).

The present study is organized in the subsequent manner. Section 2 elaborates the theoretical framework, followed by the estimation procedures and data description in Section 3. The literature arguments are presented along both sections. Estimation results are then presented and discussed in Section 4. Finally, conclusions are drawn in the closing section.

2. International Parity Conditions and Empirical Framework

PPP and RIP constitute as fundamental building block of international macroeconomics. PPP theorem requires a constant real exchange which at least exhibits reversion towards the long run mean rate over time, and not driven by stochastic trends. On the other hand, RIP is verified via the real interest differential hypothesis or real interest co-movement that support for financial asset substitutability and capital mobility across borders. If we let s_t be the log spot exchange rate, p_t^* and p_t be the log foreign and domestic price levels respectively, the PPP condition is defined as

$$s_t = p_t - p_t^* \tag{1}$$

Real exchange rates (REX), q_t (in logarithm) as deviation from the PPP is then given by

$$q_t = s_t + p_t^* - p_t \tag{2}$$

And, the ex ante PPP can be shown as

$$\Delta s_{t,t+k}^e = \pi_{t,t+k}^e - \pi_{,t+kt}^{e*} \tag{3}$$

which imply that PPP holds with expected depreciation ($\Delta s_{t,t+k}^e$) equals the expected inflation differential, and * denotes foreign variables. Subsequently, RIP can be obtained by combining the Fisher effect in each country, the ex ante PPP and the Uncovered Interest Parity (UIP) relationship. UIP anticipates expected depreciation as being explained by interest rate differentials so that

$$\Delta s_{t,t+k}^e = i_t^k - i_t^{k*} \tag{4}$$

Equating (3) and (4) thus yields $i_t^k - \pi_{t,t+k}^e = i_t^{k*} - \pi_{,t+kt}^{e*}$. If Fisher equation holds so that real interest equals nominal interest minus expected inflation, the ex ante RIP condition will be

$$E_t(r_{t+k}) = E_t(r_{t+k}^*) \tag{5}$$

When rational expectations are considered, ex post RIP also implies ex ante RIP. And, the Real Interest differential (x_t) as deviation from RIP is shown as

$$r_t - r_t^* = x_t \tag{6}$$

Given the respective specification of PPP and RIP in (2) and (6), both international parities hold if REX and RID are mean reverting. Suppose that q_t and x_t follow AR (1) process, then

$$q_t = \rho q_{t-1} + \varepsilon_t \tag{7}$$

and $\quad x_t = \varpi x_{t-1} + \mu_t \quad (8)$

where $0 < |\rho| < 1$ and $0 < |\varpi| < 1$ whereas ε_t and μ_t are white noise innovations. Evidence of long run PPP and RIP can be verified by a test of unit root in REX (q_t) and RID (x_t), say, the ADF regression with intercept and time trend which is given by

$$\Delta g_t = \mu + \beta t + \varphi g_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta g_{t-i} + \varepsilon_t \quad (9)$$

where g_t represents q_t or x_t. Δg_t is the first difference of REX or RID, k is the number of lagged g_{t-i} whilst ε_t is the error term. To be consistent with the international parities, both q_t and x_t must exhibit mean reversion behavior devoid of a unit root. The φ is to be significantly less than 0. Otherwise, deviations from PPP or RIP are permanent after shocks.

While PPP is an elegant hypothesis, early studies have shown that it fails to hold empirically (e.g. Edison, 1985; Frankel, 1986; Meese and Rogoff, 1988; Mark, 1990; Edison and Pauls, 1993). Likewise, the empirical literature does not support entirely the mean reversion behaviour of RID (see *inter alia* Mishkin, 1984; Cumby and Obstfeld, 1984; Frankel and MacArthur, 1988). The consensus arrived by recent literature survey (Rogoff, 1996; Taylor and Taylor, 2004) suggests that despite the presence of excessive short-term exchange rate volatility, the deviations from the long run equilibrium PPP rates are too persistent with the estimated half-life of real exchange shocks at about 3-5 years. For stationary REX and RID, the degree of mean reversion and extent of deviations can be further estimated by half-life, h – a concept defined as the horizon at which the percentage deviation from the long run equilibrium of PPP or RIP is one-half. By formula, $h = \frac{\ln(1/2)}{\ln(\alpha)}$, where $\varphi = (\alpha - 1)$. The two-sided 95% confidence intervals of the half-life which are based on normal sampling distributions is then defined as $\hat{h} \pm 1.96\hat{\sigma}_{\hat{\alpha}} \left(\frac{\ln(0.5)}{\hat{\alpha}} [\ln(\hat{\alpha})]^{-2} \right)$, where $\hat{\sigma}_{\hat{\alpha}}$ is an estimate of the standard deviation of α. Lately, Rossi (2005) defined half-life as $h = \ln(0.5) b(1) / \ln \alpha$ with $b(1) = (1 - \sum_{j=1}^{k} \alpha^*_{j-1})$ being the correction factor that sums the estimated AR coefficients of an AR (ρ) model fitted onto the residuals of the ADF regression. In present study, we applied both methods on REX and RID series which are found stationary.

3.1 Univariate Unit Root Test in Presence of Level Shifts

The ADF test may be distorted, however, if a potential structural break (currency crises, oil shocks, Great Crash, etc.) in the series is simply ignored (Perron, 1989). The issue was tackled in recent assessment of both theorems using various methods (Narayan, 2006; Holmes et al., 2011; Chan et al., 2011). For instance, if real exchange rates are subjected to structural breaks, then large and permanent devaluations of the currencies during a currency crisis will bias the test toward acceptance of the unit root hypothesis. Likewise, cross-border real interest may vary for the period of monetary adjustments due to hyperinflations or currency instability. Among others, Saikkonen and Lütkepohl (2002, SL hereinafter) and Lanne et al. (2002) developed break models which add to the deterministic term shift functions of a general nonlinear form using GLS de-trending procedure. The approach is extended to estimate unknown break dates by Lanne et al. (2003). Unlike much of the literature that followed dealt with the case in which a break occurs during one period only, nonlinear break tests follow the reasoning logic that breaks occur over a number of periods and display smooth transition to a new level. Say, a level shift function, which is here denoted by a general nonlinear form $f_t(\theta)'\gamma$, is added to the deterministic term, ε_t of the data generating process. Hence, the model of

$$g_t = \varepsilon_0 + \varepsilon_1 t + f_t(\theta)'\gamma + v_t \tag{10}$$

is shown, where θ and γ are unknown parameters or parameter vectors, whereas v_t are residual errors generated by an AR(p) process with possible unit root. In this study, we consider the shift function based on the exponential distribution function which allows for a nonlinear gradual shift to a new level starting at time T_B,

$$f_t(\theta) = \begin{cases} 0, & t < T_B \\ 1 - \exp\{-\theta(t - T_B + 1)\}, & t \geq T_B \end{cases}. \tag{11}$$

In the shift term $f_t(\theta)'\gamma$, both θ and γ are scalar parameters. θ is to be positive real line ($\theta > 0$), whereas γ may assumes any value. The asymptotic null distribution is nonstandard and critical values are tabulated in Lanne et al. (2002). In applying this test, one has to the AR order as well as the shift date T_B. Lanne et al. (2002) suggested that we should chose a reasonable large AR order and then pick the break date which minimized the GLS objective function used to estimate the parameters of the deterministic part.

3.2 First and Second Generation Panel Unit Root Tests

Recent studies have also progressed into panel tests of unit root and cointegration, to uncover more evidence for PPP (e.g. Wu, 1996; Papell, 1997; O'Connell, 1998; Bahrumshah, *et al.*, 2007) and RIP (e.g. Holmes, 2002; Holmes, *et al.*, 2011; Baharumshah, *et al.*, 2011). The advantages of panel tests rely on the exploitation of cross-country variations of the data and the increased in sample size, which yield higher test power in the estimation.

Among the first generation panel tests, Levin, Lin and Chu (2002, LLC) proposed to modify the ADF statistics based on homogenous pooled statistics. An estimate of the coefficient α may be obtained from proxies for Δg_{it} and g_{it} which are standardized and free of autocorrelations and deterministic components, such that:

$$\Delta\tilde{g}_{it} = \alpha\tilde{g}_{it-1} + \eta_t \tag{12}$$

where $\Delta\tilde{g}_{it} = (\Delta\bar{g}_{it} / se_i)$ and, $\tilde{g}_{it-1} = (\bar{g}_{it-1} / se_i)$, with s_i being the estimated standard error from estimating single ADF statistics of the REX and RID series . Then, LLC show that under the null, a modified t-statistics for the resulting $\hat{\alpha}$ is asymptotically normally distributed

$$t_\alpha^* = \frac{t_\alpha - (NT)S_N\hat{\alpha}^{-2}se(\hat{\alpha})\mu_{mT}^*}{\alpha_{mT}^*} \to N(0,1) \tag{13}$$

where t_α^* is the standard t-statistics for $\hat{\alpha} = 0$, $\hat{\alpha}^2$ is the estimated variance of the error term η, $se(\hat{\alpha})$ is the standard error of $\hat{\alpha}$, S_N is the mean of the ratios of the long run standard deviation to the innovation standard deviation for each individual series, which is derived using kernel-based techniques, μ^{*}_{mT} and α^{*}_{mT} are adjustment terms for the mean and standard deviation respectively, and lastly $T = T - (\sum_{i} p_i / N) - 1$.

Im, Pesaran and Shin (2003, hereafter IPS) then proposed a popular panel test that assume cross-sectional independence among panel units (except for common time effects), but allows for heterogeneity in the form of individual deterministic effects (constant and/or linear time trend), and heterogeneous serial correlation structure of the error terms. The IPS testing procedure follows the mean group approach: the *t*-bar statistics and the group mean Lagrange Multiplier test (*LM*-bar). Conceptually, the IPS test is a way of combining the evidence on the unit root hypothesis from the *N* unit tests performed on the *N* cross-section units. Through Monte Carlo experiments, the average *LM* and the *t*-statistics have better finite sample properties than the homogenous panel tests. Briefly, the test statistics are given by

$$\Gamma_{\bar{t}} = \frac{\sqrt{N}\left\{\bar{t}_{NT} - E(t_{iT} \mid \beta_i = 0)\right\}}{\sqrt{Var(t_{iT} \mid \beta_i = 0)}} \quad \Rightarrow \text{N (0,1) where } \bar{t}_{NT} = \frac{1}{N}\sum\nolimits_{i=1}^{N} t_{iT} \tag{14}$$

and

$$\Gamma_{\overline{LM}} = \frac{\sqrt{N}\left\{\overline{LM}_{NT} - E(LM_{iT} \mid \beta_i = 0)\right\}}{\sqrt{Var(LM_{iT} \mid \beta_i = 0)}} \Rightarrow \text{N (0,1) where } \overline{LM}_{NT} = \frac{1}{N}\sum_{i=1}^{N} LM_{iT} \tag{15}$$

such that $\bar{t}_{\mathrm{NT}}$ is based on averaging individual ADF tests while $\overline{LM}_{\mathrm{NT}}$ is the average across the group. Both means $E(t_{iT} \mid \beta_i = 0)$, $E(LM_{iT} \mid \beta_i = 0)$ and both variances $Var(t_{iT} \mid \beta_i = 0)$, $Var(LM_{iT} \mid \beta_i = 0)$ are obtained from the Monte Carlo simulations with i= 1,2,…,N.

The first generation panel unit root tests (LLC, IPS) discussed earlier assumes that the panel members are independent so that a Gaussian distribution can be justified by central limit arguments. In our case, this assumption can be overly restrictive because international parity conditions are expressed relative to the same benchmark as suggested in Dreger (2010) and others. In what follows the presence of cross dependencies across panel members can lead to considerable size distortions and power loss in panel tests (Banerjee et al., 2004; Pesaran, 2007; Breitung and Pesaran, 2008). While some scholars (e.g. Bai and Ng, 2004; Moon and Perron, 2004; among others) focused on the residual factor models to capture the cross-sectional dependency, Pesaran (2006) proposed that cross-sectional means of differenced data, and cross-section mean of lagged data are good proxies for unknown factors. The idea is applied in Pesaran (2007) to proxy for unobserved factors instead of using factor estimation which involves estimating the number of factors and then the factors themselves. Specifically, Pesaran (2007) proposed two types of estimation namely Common Correlated Effects Mean Group (CMG) and Common Correlated Effects Pooled (CEP). Individual specific regressors are filtered by taking the average across cross section to eliminate the effects of the unobserved common factors. The OLS procedure is taken to regress the dependent variable with regressor, the mean of cross-section dependent and individual specific regressor. Consider the dynamic panel model:

$$g_{it} = \alpha_i + \beta_i g_{it-1} + e_{it}\,,\ i = 1,2,\ldots\ldots,N,\ \ t = 1,2,\ldots\ldots,T \tag{16}$$

where α_i, β_i are parameters and differ across *i*, g_{it-1} is the first lagged value of REX or RID, and e_{it} is the random errors. In the present of cross dependency, the random errors will have the following form:

$$e_{it} = \gamma_i f_t + \omega_{it}, \; i = 1,2,\ldots\ldots,N, \;\; t = 1,2,\ldots\ldots,T \qquad (17)$$

where f_t is the latent factors, γ_i are factors loadings that probably influenced by the factors and ω_{it} is the random errors of e_{it}. Following Pesaran (2007), two assumptions will be considered before testing for a unit root in panel model: (i) the ω_{it} and f_t are serially uncorrelated for each *i* with zero mean and the variance, $0 < \sigma_i^2 < \infty$, and (ii) the ω_{it}, f_t and γ_i are independently distributed for all *i*. Eq. (16) subtracted with g_{it-1}:

$$\Delta g_{it} = \alpha_i + b_i g_{it-1} + \gamma_i f_t + \omega_{it}, \; i = 1,2,\ldots\ldots,N, \;\; t = 1,2,\ldots\ldots,T \qquad (18)$$

where $\Delta g_{it} = g_{it} - g_{it-1}$ and $b_i = \beta_i - 1$. The OLS estimate for b_i is based on the regression:

$$\Delta g_{it} = \alpha_i + b_i g_{it-1} + c_i \Delta \bar{g}_t + d_i \bar{g}_{t-1} + \omega_{it}, \; i = 1,2,\ldots\ldots,N, \;\; t = 1,2,\ldots\ldots,T \qquad (19)$$

Under the null the model is unit root ($b_i = 0$ for all *i*) against stationary ($b_i < 0$ for some *i*), the test statistics, $t = \dfrac{\sqrt{TN}(\bar{b} - b)}{\sqrt{\sum_{i=1}^{N} Var(b_i)}} \xrightarrow{d} N(0,1)$ with $t = \dfrac{\sqrt{T}(\hat{b}_i - b)}{\sqrt{Var(b_i)}} \xrightarrow{d} N(0,1)$ where

$$b_i = \left(g_{i,-1}^T \bar{M} g_{i,-1}^T\right)^{-1}\left(g_{i,-1}^T \bar{M} \Delta g_i\right), \mathrm{var}(b_i) = \hat{\sigma}_i^2 \left[g_{i,-1}^T \bar{M} g_{i,-1}^T / T\right]^{-1}, \hat{\sigma}_i^2 = \frac{(\Delta g_i - g_{i,-1}\hat{b}_i)^T \bar{M}((\Delta g_i - g_{i,-1}\hat{b}_i)}{T-4}$$

and, the properties of $\bar{b} = \sum_{i=1}^{N} b_i \Big/ N$, $\dfrac{\sqrt{TN}(\bar{b} - b)}{\sqrt{\sum_{i=1}^{N} Var(b_i)}} \xrightarrow{d} N(0,1)$.

However, the properties of CMG is $\dfrac{\sqrt{N}(\hat{\beta}_{CMG} - \beta)}{\sqrt{\Sigma_{CMG}}} \rightarrow N(0,1)$ where $\hat{\beta}_{CMG} = \dfrac{\sum_{i=1}^{N} \hat{\beta}_i}{N}$ and

$\hat{\Sigma}_{CMG} = \dfrac{\sum_{i=1}^{N} (\hat{\beta}_i - \hat{\beta}_{CMG})(\hat{\beta}_i - \hat{\beta}_{CMG})^T}{N-1}$. The $\hat{\beta}_i$ for $i = 1,2,\ldots\ldots,N$ are obtained by computing $\hat{\beta}_i = (X_i^T \bar{M} X_i)^{-1} X_i^T \bar{M} g_i$ and $\bar{M}$ is defined as $\bar{M} = \mathrm{I}_t - \bar{H}\left(\bar{H}^T \bar{H}\right)^{-1} \bar{H}^T$ with $\bar{H} = (D, \bar{g}_t, \bar{g}_{t-1})$. I_t is a unit matrix of order *T*x*T* and $\bar{H}$ is the combinations of dummy variables, average of cross section of the first difference of g_{it} and its g_{it-1}. The properties of CEP is $\dfrac{\sqrt{N}(\hat{\beta}_{CEP} - \beta)}{\sqrt{\Sigma_{CEP}}} \rightarrow N(0,1)$ where $\hat{\beta}_{CEP} = (\sum_{i=1}^{N} X_i^T \bar{M} X_i)^{-1} (\sum_{i=1}^{N} X_i^T \bar{M} g X_i^T)$ and $\hat{\Sigma}_{CEP} = \hat{\psi}^{-1} \hat{V} \hat{\psi}^{-1}$.

The $\hat{\psi} = \dfrac{\sum_{i=1}^{N} X_i^T \bar{M} X_i}{NT}$ and $\hat{V} = \dfrac{\sum_{i=1}^{N} (X_i^T \bar{M} X_i)(\hat{\beta}_i - \hat{\beta}_{CMG})(\hat{\beta}_i - \hat{\beta}_{CMG})^T (X_i^T \bar{M} X_i)^T}{(N-1)T^2}$.

3.3 Data Description

Various tests outlined in the previous section are applied to a sample of monthly observation for China and her thirteen major trading partners in the Asia Pacific. Except India, all trading partners are APEC members including the economic giants (US, Japan), the Oceania economies (Australia, New Zealand), the developed NIE-4 (Hong Kong SAR, Singapore, South Korea, Taiwan) and the developing ASEAN-4 (Indonesia, Malaysia, Philippines, Thailand). Our joint investigation of PPP and RIP involves the bilateral real exchange rates (REX) and real interest rate differentials (RID) of China-APEC. The construction of thirteen China-denominated REX is based on equation (2), which consists of nominal Yuan-based exchanges rates, individual APEC CPI as domestic price and China CPI as foreign price. As for RID, China is again considered as foreign country (numeraire) and we follow the Fisher equation to construct real interest rate by subtracting the expected inflation from nominal interest rate. Since ex post RIP implies ex ante RIP, expected inflation is proxy by actual inflation. The nominal interest rates used in the study are generally non-control and medium term lending rates due to the fact that long-term interest rates, such as government bond yields are incomplete or unavailable for most of these Asian countries. To uphold the consistency and reliability of the data, we cross check with various data sources namely Datastream, International Financial Statistics of IMF, and Central Banks of respective economies.

4.1 Empirical Discussion of Endogenous Breaks and Unit Root Tests

It is widely recognized that classical unit root tests might be biased by the presence of structural breaks and nonlinearities in the deterministic components. An alternative approach that captures the structural breaks with a smoother functional form for the transition period could be more informative. For this purpose, we apply the SL test with the optimal lag length (k) being determined by the standard Schwarz Information Criterion (SIC). As can be seen in Table 1, all the exponential shift parameters appear to be highly significant to capture the endogenous shift dates. For REX, endogenous break(s) occur mainly in 1993, except for two where the break date is detected in 1997/98. The first break date is due to the major downward adjustment (appreciation) of Chinese Yuan in 1993/94 against the USD and other major currencies. The second break date coincides with the Asia financial turmoil that witnessed a sharp fall of the East Asian currencies. As can be seen in the table, only four out of 13 Yuan-based REX (Taiwan, Indonesia, India and New Zealand) rejected the unit root null hypothesis at the indicated significant levels. Results based on the SL tests indicate the absence of mean reversion behaviors even when graduate shifts are allowed in the model. If this is true, then for any shocks on the REX series, deviations will be too persistent to witness necessary adjustment to the equilibrium level and the PPP puzzle remains unsolved. Such finding is inconsistent with the recent USD- and Japanese Yen-based PPP studies, but tend to support the argument that Chinese Yuan is misaligned and inconsistent with the PPP rules. [**Insert Table 1]**

As for RIDs, most breaks occurred at 1988 when China experienced high inflation that resulted in imbalance rate of real interest. Also, some adjustments of interest rates were found in 1998/99 among the crisis-affected nations to defend their currencies and to tackle the stagflation (e.g. Indonesia). Unlike the results from the PPP presented above, most China denominated-RID (except South Korea) have exhibited mean reversion behavior and lend support for RIP. Nevertheless, the rejection of univariate unit root alone is necessary but neither sufficient to gauge the degree of mean reversion of APEC-China series as a group nor to identify the potential changes in the process of integration due to market and policy reforms over time. Next, we proceed with the panel tests that utilize both cross-sectional and time series information to allow us to validate the respective PPP and RIP condition using sub-sample analysis.

Information about the endogenous break dates has been useful to construct our sub-samples in panels. Considering the frequency of break dates, we separate the study periods into 1987-1993, 1994-2007, 1987-1997, 1998-2007 and 1987-2007. However, we are unable to consider the 1988 break as the sample size is too short and inappropriate for econometric estimation. To improve the robustness of our findings, the homogenous and heterogeneous panel tests are both conducted. For the early sub-periods of 1987-1993 and 1987-1997, the panel results of REX support the SL findings reported earlier which generally against the PPP, suggesting the inflexibility of exchange rate and deviations from equilibrium rate are permanent. This is indeed the period when Chinese Yuan practiced multiple rates and the official rates were de facto crawling band around USD (+/- 2%) with the premium peaks at 124% on June 1991.

Even when the full sample size is considered, null hypothesis of unit root fail to be rejected and no evidence of mean reversion is captured. The results differ and improved drastically when the sub-sample of post-liberalization (1994-2007) and post-crisis (1998-2007) are considered. Rejections of unit roots are highly significant as reported by LLC and IPS tests, implying that the deviations of the group of Yuan-based REX are now temporal, and exchange rates are more responsive to changes in price ratios. These are mainly attributed to the unification of China's two main currency rates in 1994 and the deregulation on foreign invested enterprises in exchanging funds freely at selected banks without approval from the State Administration for Exchange Control (SAEC) in 1996 that drive the RMB a step further towards the full convertibility (see Zhang, 1999). The adjustment of under-valued Renminbi (RMB) since 2005 may also en route for some extent of market completeness by PPP rules. But overall, the liberalization process is still insufficient to display full support for PPP and further flexibility in the exchange rate regime is needed. **[Insert table 2]**

A somewhat comparable trend of mean reversion behavior is found when the APEC-China real interest differentials (RID) are taken as a group. For instance, LLC has failed to reject the null hypothesis of common unit root for two early sub-samples but highly rejected for two late sub-samples. Similar but not identical, the IPS heterogeneous panel test detected weak rejection of individual unit roots for the early period sub-samples but strong rejection of unit roots for late period sub-samples. Putting them together, the supports for RIP are general weaker during pre-liberalization era but improved evidently for the post-liberalization period, before and after the crisis (Table 2). In most cases, both the univariate and panel tests of unit root seem more supportive of RIP rather than PPP for China vis-à-vis Asia Pacific economies.

Nevertheless, final conclusion is yet to be drawn at this stage[5]. There are still questions if the sub-sample analysis bias toward unit root null or alternative when the cross sectional dependency present in the series. The Lagrange Multiplier of Bruesch and Pagan (1980)'s and Pesaran (2007)'s cross dependency tests are, by this means, deployed for additional analysis. The result is reported in Table 3. Column 2 and column 7 show the respective sample pair wise correlation of the residuals ($\left|\hat{\rho}_{ij}\right|$) for REX-China and RID-China. CD_{lm} refers to Lagrange Multiplier of Bruesch and Pagan and PCD refers to Pesaran's cross sectional dependence tests. Under the null of no cross dependency, both of the tests overwhelming reject the null in favor to there is at least one cross sectional dependence at 5 % significance level in all the sub-sample for REX and RID. These have prompted us to utilize the Pesaran's CMG and CEP panel unit root tests to account for cross sectional dependence for the sub-sample panel. As shown in columns 5 and 10 for CMG and columns 6 and 11 for CEP, both the RIP and PPP hold significantly in all but one case. We found the REX is nonstationary and against the PPP over the 1987M1-1997M12. Our findings are consistent with Narayan (2006) who found stationarity with breaks of India's bilateral exchange rate vis-a-vis fifteen out of sixteen of its major trading partner. **[Insert Table 3]**

The empirical evidence, as overall, coincides with the financial liberalization process and the gradual ruling out of restrictions on capital movements in APEC, including China. In June 1996, the ceiling rates of inter-bank loans were removed and the interest rates have expanded twice in China within 1998-99 while state-owned financial institutions are allowed to be commercialized. By September 2000, the controls on large fixed deposits and foreign currency loans were lifted and the China Banking Association took over the responsibility of interest rates decision on small foreign currency deposit. Because China is taken as base country, support for RIP would confirm the improved influence of China in the regional capital markets since 1990s. Future fluctuations of the APEC real interest rates can possibly be determined or forecasted, using the Chinese real rates as part of the information set. In addition, the results do indicate the benefits of using panel tests in exploiting the cross cross-country variations of the data, thus, yielding higher test power in the sub-sample and also whole sample estimation over time.

4.2 Half-Life Estimation and Confidence Intervals

To obtain an insight into the degree of mean reversion of REX and RID as further justification of PPP and RIP, the estimation of half-life for series that are found stationary is essential. But since the point estimates of half-life may provide an incomplete picture of the speed of convergence towards the equilibrium rates in long run, the corresponding confidence intervals are also computed. Such practice offers better indications of the uncertainty around the estimates of half-life. For univariate series, this study estimates the half-life based on the AR (ρ) method and the correction factor model proposed by Rossi (2005). For panel series with sub-samples, only the AR (ρ) method is employed.

[5] The outcome of the first generation panel tests is sensitive to the selection of series included in the group, as the null hypothesis of a common unit root (homogenous) may be rejected even if only one of the series is stationary. As a result, several studies proceed with the heterogeneous panel tests (allowed for cross-sectional independence) with uncorrelated errors or the second generation panel tests that account for cross-section correlation of errors (see Breitung and Pesaran, 2008).

We first select four REX series that support PPP for estimation and 16-59 months (1.3-4.9 years) of half-life is reported by the classical method (Table 4). All except Taiwan has reported slightly shorter half-life when the Rossi method is applied and displayed moderate speed of adjustments to the equilibrium PPP rate. In the panel analysis with all APEC-13 pooled as a group, the post-liberalization and post-crisis period recorded half-life around 18-24 months (1.5-2 years). The standard errors are considered miniature and contribute to a less widen but stable confident intervals. There are signs that deviations of REX exhibit somewhat faster adjustments back to the long run PPP since 1994. **[Insert Table 4]**

On the other hand, supports for RIP as indication of financial integration are somewhat greater than supports for PPP. Univariate series averagely show 14 – 28 months (1.2–2.3 years) of half-life. Then again, the scale of half-life drops to about 3 – 11.3 months under the Rossi estimation. For panel analysis, full sample (1987-2007) half-life is approximately 31 months. As for the post-liberalization with (1998-2007) and without the crisis (1994-2007), the half-lives are recorded at 8 and 27 months respectively. Consistent with the panel results, the shortened half-life bounded with more stable confident intervals has provided solid evidence in support for the RIP among APEC-China. The signs of decreasing deviations from RIP are evident and in line with the increased regional financial integration prompted by financial liberalization, technological breakthroughs, and growth in the volume of trade in recent years (Baharumshah, *et al.*, 2011).

All in all we find that the speed of mean reversion is high, indicating that RIDs tend to be short-lived. Allowing for the possibility of structural breaks, we find even shorter-lived deviation from equilibrium. This evidence is supportive of high degree of market integration, which is consistent with financial liberalization and the emergence of global financial markets. The varying speed of the adjustments to long run PPP and RIP across the countries reviewed may reflects China's position in pursuing liberalization in good and capital markets at multi-speed. The rapid growth in the regional capital flows has contributed to cross-border investments and optimal allocation of resources and, in some cases has facilitated the movement towards financial convergence and closer monetary cooperation. Conservative policies directed at increasing domestic savings to increase the rate of capital formation and hence productivity growth, are no longer the solely option in open economy macroeconomics. Instead, cross-border capital flows raise the chances of risk-sharing, portfolio diversification, and thus enable countries in the Asia Pacific region to smooth out consumption.

5. Conclusion and Policy Implications

This paper conducts a joint investigation of two international parities, namely the PPP and RIP, to assess the extent of goods and capital market integration between China and her 13 trading partners in Asia Pacific region. Endogenous and exponential breaks are confirmed for the real exchange and real interest differential series, which mostly occur in 1988, 1993/94 and 1997/98. The break dates coincides with the major events in the region. Our major findings are three-fold. First, we observe that RIP holds better than PPP, suggesting the greater financial integration than trade integration among APEC-China. Second, both parities tend to hold better as one move to the recent years. Third, China and APEC has improved the ability to absorb regional shocks as indicated by the shortened half-life reported over time, especially when the post-Asia crisis era is included.

Putting together, the greater integration among APEC-China implies the better equalization of the marginal utility of home and foreign currency (Renmimbi), which in turn allows for better risk sharing. The integration process is attributed not only to the liberalization process among the APEC economies, but also to the Chinese trade policy and the regional commitment for the ASEAN+3+2+1 cooperation. Besides, the prospect of WTO membership is indeed instrumental for China to move towards liberalizing its external sectors and capital accounts. This coincides with our finding of mean reversion behavior in the China-based real interest differentials, which implies the increased influence of Chinese investments in the regional capital market. Moreover, the shorter half-lives reported over time encourage us to foresee a brighter feasibility towards regional financial deepening and regional currency arrangements that anchored by China. By taking cooperative action, China and APEC members would be in a better position to resist the adverse consequences of sudden and sizeable movements in global capital, and the potentially deleterious effects that may decelerate the growth and development of domestic economies. After all, monetary and exchange rate policy cooperation in East Asia would enable this region to exert an important influence upon the future evolution of the global trade and financial system.

It is important to note, however, that RIP holds better than PPP may also raise some concerns on the sequencing issue of economic integration. PPP does not hold fully and the Yuan-denominated currencies are still not highly competent by the PPP rules. China's market size and its role as a production hub are yet sufficient to draw a full support for PPP as indication of perfect trade integration among APEC-China. Or, in other words, regional trade competition is yet sufficient to eliminate prices arbitrage to reflect the exchange value of Chinese Yuan. While the more liberalized exchange rate regimes among APEC members may have facilitated for better integration, the prolonged undervalued Renmimbi, has as well exerted some drawbacks in the PPP theorem especially during 1980s-1990s. Further flexibility in the Chinese exchange rate regime is expected.

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Table 1: Univariate Unit Root Test with Endogenous Break

	REX –CHINA			RID –CHINA		
	k	Break	SL test	*k*	Break	SL test
US	7	1993M6	-2.515	3	1988M8	-2.920 [b]
Japan	1	1993M6	-1.764	5	1988M8	-2.951 [b]
India	4	1993M6	-2.707 [a]	5	1998M11	-2.998 [b]
Australia	1	1993M7	-1.984	3	1990M1	-2.747 [a]
New Zealand	7	1993M6	-2.614 [a]	2	1989M8	-3.191 [b]
Hong Kong	1	1993M6	-2.300	1	1988M8	-2.880 [b]
Taiwan	2	1993M7	-2.762 [a]	6	1989M9	-3.284 [b]
South Korea	2	1997M12	-2.062	2	1988M10	-2.312
Singapore	5	1993M6	-1.480	2	1992M4	-2.906 [b]
Indonesia	6	1998M1	-2.648 [a]	4	1999M2	-2.963 [b]
Malaysia	7	1993M6	-2.495	5	1988M8	-3.012 [b]
Philippines	1	1993M6	-1.562	2	1991M11	-3.221 [b]
Thailand	5	1993M6	-1.647	3	1989M8	-2.963 [b]

Critical values

1% [c]	5% [b]	10% [a]
-3.48	-2.88	-2.58

Notes: ([a]), ([b]) and ([c]) denote for the significant level at 10%, 5% and 1% respectively. Critical values are obtained from Lanne, Lütkepohl, and Saikkonen (2002).

Table 2: First Generation Panel Unit Root Tests

	LLC- Homogeneous Panel Test		IPS-Heterogeneous Panel Test	
	REX-CHINA	**RID-CHINA**	**REX -CHINA**	**RID –CHINA**
1987M1-1993M12	2.887 (0.998)	-0.861 (0.195)	1.158 (0.876)	-1.595 [a] (0.055)
1987M1-1997M12	1.428 (0.923)	-0.155 (0.438)	2.476 (0.993)	-2.465 [b] (0.007)
1987M1-2007M1	-0.537 (0.296)	-2.065 [b] (0.020)	-0.376 (0.354)	-2.962 [c] (0.002)
1994M1-2007M1	-6.616 [c] (0.000)	-2.727 [c] (0.003)	-5.056 [c] (0.000)	-2.367 [c] (0.009)
1998M1-2007M1	-2.040 [b] (0.021)	-2.676 [c] (0.004)	-2.439 [c] (0.007)	-4.804 [c] (0.00)

Note: ([a]), ([b]) and ([c]) denote for the significant level at 10%, 5% and 1% respectively.

Table 3: Second Generation Panel Unit Root Tests

	REX-CHINA					RID-CHINA				
	$\lvert\hat{\rho}_{ij}\rvert$	CD_{lm}	PCD	CMG	CEP	$\lvert\hat{\rho}_{ij}\rvert$	CD_{lm}	PCD	CMG	CEP
1987M1-1993M12	0.880	5044.471^{c} (0.000)	70.782^{c} (0.000)	-4.394^{c} (0.000)	-1.709^{b} (0.044)	0.787	4033.751^{c} (0.000)	62.960^{c} (0.000)	-6.196^{c} (0.000)	-4.664^{c} (0.000)
1987M1-1997M12	0.789	6371.780^{c} (0.000)	79.403^{c} (0.000)	0.094 (0.462)	-1.011 (0.156)	0.782	6308.004^{c} (0.000)	78.722^{c} (0.000)	-4.165^{c} (0.000)	-6.165^{c} (0.000)
1987M1- 2007M1	0.693	9239.572^{c} (0.000)	94.608^{c} (0.000)	-6.126^{c} (0.000)	-3.702^{c} (0.000)	0.668	8974.175^{c} (0.000)	91.233^{c} (0.000)	-6.005^{c} (0.000)	-2.699^{c} (0.003)
1994M1- 2007M1	0.525	3589.059^{c} (0.000)	57.729^{c} (0.000)	-4.634^{c} (0.000)	-2.882^{c} (0.002)	0.566	4482.899^{c} (0.000)	62.213^{c} (0.000)	-6.515^{c} (0.000)	-3.221^{c} (0.001)
1998M1- 2007M1	0.541	2614.405^{c} (0.000)	49.426^{c} (0.000)	-5.150^{c} (0.000)	-3.770^{c} (0.000)	0.478	2330.419^{c} (0.000)	43.504^{c} (0.000)	-6.706^{c} (0.000)	-10.689^{c} (0.000)

Notes:

(1) (^{b}) and (^{c}) denote significant at 5% and 1% significant levels, respectively.

(2) $\lvert\hat{\rho}_{ij}\rvert$ denotes the sample wise correlation of the residual denoted as $\hat{\rho}_{ij} = \hat{\rho}_{jj} = \sum_{t=1}^{T} \hat{e}_{it}\hat{e}_{jt} \Big/ \left(\sum_{t=1}^{T} \hat{e}_{it}^{2}\right)^{1/2} \left(\sum_{t=1}^{T} \hat{e}_{jt}^{2}\right)^{1/2}$.

(3) Reject H_0 when $CD_{lm} = T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij} > \chi^{2}_{(N(N-1)/2)} = 99.62$ and $PCD = \sqrt{2T/N(N-1)}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij} > N(0,1) = 1.96$.

Table 4: Univariate Half-life Estimations

	REX-CHINA		RID-CHINA	
	HL- AR(ρ) [95%CI]	Rossi (2005) [95%CI]	HL- AR(ρ) [95%CI]	Rossi (2005) [95%CI]
US	-	-	27.33 [9.28, 45.38]	8.08 [0, 26.13]
Japan	-	-	24.85 [8.43, 41.26]	7.24 [0, 23.65]
India	15.87 [3.91, 27.84]	14.97 [3.01, 26.94]	19.16 [5.43, 32.90]	8.52 [0, 22.25]
Australia	-	-	28.05 [6.19, 49.91]	11.29 [0, 33.15]
New Zealand	26.13 [1.29, 50.96]	16.59 [0, 41.43]	25.31 [9.32, 41.31]	7.96 [0, 23.96]
Hong Kong	-	-	21.13 [5.49,36.77]	11.27 [0, 26.91]
Taiwan	58.85 [0, 160.04]	61.18 [0, 162.38]	15.32 [5.39,25.25]	6.20 [0, 16.14]
South Korea	-	-	-	-
Singapore	-	-	24.38 [8.23, 40.54]	8.53 [0, 24.69]
Indonesia	19.67 [2.92, 36.42]	12.89 [0, 29.64]	13.98 [7.55, 20.42]	3.06 [0, 9.50]
Malaysia	-	-	24.48 [7.98, 40.98]	7.50 [0, 24.00]
Philippines	-	-	17.37 [5.57, 29.17]	8.92 [0, 20.71]
Thailand	-	-	25.52 [7.56, 43.48]	9.26 [0, 27.22]

Notes: Half-life is computed only for stationary series confirmed by SL test.

Table 5: Panel Half-life Estimations

	REX-CHINA		RID-CHINA	
	N	HL- AR(ρ) [95%CI]	N	HL- AR(ρ) [95%CI]
1987M1-1993M12	-	-	-	-
1987M1-1997M12	-	-	-	-
1987M1-2007M1	-	-	3108	30.96 [23.31, 38.61]
1994M1-2007M1	2041	18.10 [14.41, 21.80]	2041	27.41 [20.07, 34.75]
1998M1-2007M1	1417	23.98 [14.04, 33.92]	1417	7.60 [5.87, 9.33]

Notes: N represents the number of observations utilized in the panel analysis. Half-life is computed based on the AR (ρ) methodology only for stationary series confirmed by both LLC and IPS tests.