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A STRUCTURAL MODELING OF EXCHANGE RATE, PRICES AND INTEREST RATES BETWEEN MALAYSIA-CHINA IN THE LIBERALIZATION ERA

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Abstract
This study constructs a structural system that allows for possible interactions between the goods and capital markets for Malaysia vis-à-vis China in the liberalization era (1994: Jan to 2011: June). It encompasses the joint hypothesis of Purchasing Power (PPP) and Interest Rate Parity (IRP) conditions in the presence of $I(1)$ exogenous variables. Advanced econometric procedures including the structural VARX, VECX*, over-identifying restrictions, bootstrapping, persistent profiles and generalized variance decomposition are utilized in the analyses. The finding upholds support for both PPP and IRP, when exchange rate regime and structural breaks of Asia crisis and subprime crisis are taken into accounts. Despite the direct imported inflation, exchange rate also plays a significant role in the price transmission mechanism. And, Malaysian maintains the relative monetary autonomy against China in short run, but the price channel will affect the extent of IRP condition. Lastly, the faster pace of adjustment towards price instead of interest rate equilibrium implies the nonappearance of sequencing problem in market integration. Putting together, our model contributes as an early warning system for Malaysia's economic defense against global shocks.

Keywords: International Parity Conditions, Market Integration, Price Transmission Mechanism, Structural VARX, Bootstrapping

JEL Classification: C51, F36, F42

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1. Introduction

In the past decades, Malaysia has been closely linked to the US and Japan. But since 2009, China became Malaysia's largest trading partner—the largest source of imports and second largest export destination. Malaysia-China trade reached \$59 billion—about 18.9% of Malaysia's global trade volume, surpassing the Malaysia-US trade share (10.9%). In recent years, local banks have also introduced Renminbi Trade Settlement Services. Together, the trade and investment expansion is likely to accelerate with the formalization of a bilateral trade liberalization pact on track under the ASEAN-China Free Trade Agreement (Okamoto, 2005). While Malaysia-China economic integration has grown in greater and faster pace, there are worries that such linkage may be destructive. McKibbin and Woo (2003), for instance, suggests that the full integration of Chinese labor force into the international division of labor could de-industrialize the ASEAN (including Malaysia) when it leads to reduction of FDI flows to them. Some observers have also, directly or indirectly, related the resurgence of China since the late-1980s and the devaluation of the renminbi (or, Chinese yuan) in 1994 to the Asia financial crisis (Makin, 1997; Corsetti, *et al.*, 1999; among others).[1]

Due to the fact that China has been the major source of imports—both consumer goods and industry inputs, changes in Chinese labor costs and producer prices are highly concerned. Similarly, the increased financial risks of the Chinese asset market and their recent speculative capital flows to Malaysia have gained attention of domestic policy makers. Both Malaysia and China have maintained an undervalued exchange rate regime[2] since 1990s and the ringgit and yuan have moved closely during 1998-2005 (Chan and Hooy, 2011). Though claimed as managed float by Chinese authority, the Chinese yuan was de facto pegged to the USD at RMB8.28 from 1998 through June 2005 (Yongjian, *et al.*, 2009). Malaysia, on the hand, was officially pegged to USD at RM3.80 in similar period. Such policy coordination would imply that the chances of contagious-financial turmoil and -inflation are highly feasible between the two nations, as long as monetary sovereignty against China remained.[3] Nevertheless, the potential impacts are still questionable. Unless a comprehensive study is conducted, the transmission mechanism cannot be fully understood.

To tackle the mentioned issues, an inclusive inspection of the international parity conditions is necessary. As theoretical propositions, purchasing power parity (PPP) and interest rate parity (IRP) provide clues of how price and monetary effects are transmitted globally. By implication, PPP acts as a backward adjustment mechanism in the goods market whereas the IRP (e.g. Uncovered Interest Parity-UIP) can be thought of an arbitrage relationship that function as forward-looking market clearing mechanism in capital market (Juselius, 1995). Both theories are also popular in the assessment of goods and capital market integration (Cheung, *et al.*, 2003; Cavoli, *et al.*, 2004; Kargbo, 2009). Nevertheless, the respective empirical evidence of PPP and UIP, which has hitherto been abundant, is still inconclusive (see Rogoff, 1996; Alper, *et al.*, 2009; for recent surveys). Among China 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

[1] The fall of the Chinese yuan implied a real exchange rate appreciation for the dollar-pegged currencies in East Asia, which their fragile financial systems were unable to absorb. Some of them were thrown into prolonged current account deficits and forced to devalue their currencies in order to regain their export market share, which eventually led to the Asia financial crisis in 1997.

[2] Big Mac Indexes show that Chinese yuan and Malaysian ringgit continue to be substantially undervalued as much as 40%-45% and 25%-30% respectively, in 2010 (The Economist, *various issues*).

[3] According to the macroeconomic trinity, it is impossible for a nation to have a fixed exchange rate, free capital flows and independent monetary system at the same time.

– only for the real effective yuan but not for the real yuan/USD rates. Cheung, *et al.* (2003), in addition, examined three parity conditions consecutively and concluded that parities hold among China-Taiwan-Hong Kong. Meanwhile, Cavoli, *et al.* (2004) examined the parity conditions for ASEAN5, East Asia and China but failed to find clear indication of intensified regional financial integration. Other than the methodological concerns, a rather mixed and puzzling evidence that have accumulated on time series properties of UIP and PPP could be due to the failure account for the interdependence of adjustments in the international asset and commodity markets (Juselius, 1995; Özmen and Gökcan, 2004). The policy arguments recently extend from the validity of parity conditions to the exploration of connection and sequence between trade and financial integration among Asian members (Pomfret, 2005; Eichengreen, 2006).

With an intention to solve the mentioned problems, this paper hereby constructs the joint assessment of PPP and IRP between Malaysia-China using the structural modeling method. The study period spans from 1994: Jan to 2011: June, where both Malaysia and China are experiencing trade expansion and economic liberalization. Also, the fixed exchange rate regime is taken into account. More important, unlike previous works that study the PPP or IRP separately, we assess the interaction and transmission effects of prices, interest rates and exchange rates within a full system framework, as inspired by Johansen and Juselius (1992), Juselius (1995) and Juselius and MacDonald (2004)[4]. The modeling approach allows for the possible interactions between goods and capital market, which will potentially constitute the foundation of an early warning system particularly for Malaysia, against external shocks. Our approach also recognizes the importance of distinguishing the short-run and long-run effects in the model so that the error correction terms of the PPP and IRP are empirically valid and in line with theoretical prediction.

Before we proceed with the analysis, there are few significant considerations that distinguish our study from the literature. The first concerns the fact that Malaysia is a small and open economy. When compared to the Chinese population of 1.3 billion people, the Malaysian market size is relatively small, with only 28 million residents. Though Malaysian trade openness is now among the highest in the world (about 200% of its GDP), the economic size and financial influence are significantly lesser as compared to China. Apart from being the largest economy body in Asia (second world largest) since 2008, China has also become the world's second-largest trading nation after the US. It is thus necessary, in the methodological sense, to develop an econometric model that allows the possibility of drawing a distinction between endogenous and exogenous variables, which are integrated of I*(1)*. This paper employs the structural modeling procedures advanced by Pesaran *et al*. (2000) and Assenmacher-Wesche & Pesaran (2009). We construct a cointegrating VARX with two long-run equilibrium relationships (PPP and IRP) in the presence of I*(1)* weak exogenous or long-run forcing variables (which, in our case, the Chinese variables). A reduced-form error correction of the VECX* short-run model can then be estimated, where variables are separated into the conditional model and marginal model, respectively. Such structural modeling methodology builds on transparent and theoretically coherent foundation that offers

[4] Johansen and Juselius (1992) and Juselius (1995) argued that previous studies on international parity conditions may have overlooked the links between goods and asset markets, and partly due to the lack of a precise specification of the sampling distribution of the data. They are able to show supportive evidence for the PPP and IRP relations in the UK case when a systemic multivariate cointegration framework is adopted. Similar analyses have been performed on different series of developed nations (e.g. Australia, German, Norway, Sweden) and some non-identical but similar conclusions were observed (see *inter alia*, Sjoo, 1995; Caporale, *et al*., 1995; Juselius and MacDonald, 2004).

a practical approach to relationships suggested by economic theory. To further assess the effect of system-wide shocks on the cointegrating relations, we apply the Persistence Profile analysis developed by Pesaran and Shin (1996). Subsequently, we also gauge the out-sample causality effects using the generalized forecast error Variance Decompositions (VDCs).

Then, what follows involves the estimation issue for small sample study, particularly, in regard to the size and power properties of time series analysis. In our case, the study period covers 17 years with 210 monthly observations, which is considered short for international parities study. Given this, we use the nonparametric bootstrap method, an alternative to the large sample data tests based on asymptotic theory. Bootstrap's ability to provide asymptotic refinements often leads to a reduction of size distortions in finite sample bias and it generally yields consistent estimators and test statistics (Mantalos and Shukur, 1998; Chang, Park and Song, 2006). This method is employed to test the number of VARX cointegrating ranks. It is later applied in the estimation of log-likelihood ratio (LR) critical values for the PPP and IRP normalized (exactly identified) and over-identified restrictions as well as for the marginal model and conditional model in the VECX* error correction representation. Bootstrapping is also used to estimate the confidence intervals of Persistent Profile. Then again, the 1990s-2000s are well known as a period of financial instability and currency crises. A preliminary Zivot-Andrews (1992)'s test of endogenous break(s) on each series is conducted and we impose the break dates (e.g. Asia crisis, fix exchange rate regime, Subprime crisis) as dummy variables in the VARX and VECX* models.

Our study is organized in the following manner. Section 2 shows the theoretical representation of PPP and IRP that forms the basis of our empirical model. This is then followed by the estimation procedures of VARX and VECX* and data description. Estimation results are discussed in section 3. Finally, in section 4, conclusion and policy implications are drawn.

2. Theory and Methodology

Being the first equilibrium theory of exchange rate, the theoretical motivation for PPP is based on the assumption that internationally produced goods are perfect substitutes for domestic goods. On the other hand, the second equilibrium theory of exchange rate—UIP, states that the interest rate differential between two countries is equal to the expected change in the spot exchange rates. UIP assumes zero risk premium so that financial assets are substitutes in cross-border capital markets. If we let EX_{Mt} be the log spot exchange rate of RM/yuan, P_{Mt} and P_{Ct} be the log domestic (Malaysia) and foreign (China) price levels respectively, the PPP condition is defined as

$$P_{Mt} = P_{Ct} + EX_{Mt} \tag{1}$$

while UIP condition is represented by

$$R_{Mt} = R_{Ct} + E_t(EX_{Mt+1}) - EX_{Mt} \tag{2}$$

with R_{Mt} and R_{Ct} being the respective nominal interest rates denominated in domestic and foreign currencies compounded over the time period t - $(t$ - $1)$, and E_t (.) denotes the expected value formed at time t. When the forecast horizon grows, it seems reasonable to expect deviations from long-run PPP to be increasingly important in the formation of expectations, thereby providing a link between the goods and the capital markets. More specifically, if the expected exchange rate is given by

$$E_t(EX_{Mt+1}) = P_{Mt} - P_{Ct} \tag{3}$$

a relation combining the PPP and the UIP conditions can be derived by inserting (3) into (2):

$$R_{Mt} - R_{Ct} = P_{Mt} - P_{Ct} - EX_{Mt+1} \tag{4}$$

(l) - (4) are simple economic hypotheses which define 'long-run' equilibrium in the capital and goods markets in a very simplified world. For empirical analysis purpose, Eq. (4) will be adopted in our VARX ad VECX* estimations.

2.1 The VARX and VECX Estimation*

Pesaran *et al*. (2000) modified and generalized the approach to the problem of estimation and hypothesis testing in the context of the augmented vector error correction model. Garratt *et al*. (2003, 2006) extended the idea and developed the VECX* model along the same lines. They distinguish between an $m_y \times 1$ vector of endogenous variables y_t and an $m_x \times 1$ vector of exogenous I(1) variables x_t among the core variables in $z_t = (y'_t, x'_t)$ with $m = m_y + m_x$. In our case, the two exogenous variable as 'long-run forcing' variables are the Chinese price and interest rates. 'Forcing' variable means that changes in P_{Ct} and R_{Ct} have a direct influence on, but not affected by Malaysian variables in the model. This ends up with a conditional vector error correction model (VECX*) with five variables and two structural cointegration relations, in which the two long-run relations (r = 2) correspond to PPP and IRP.

Since our study covers the period of the Asia financial crisis, fixed exchange regime and the subprime crisis, structural break(s) are necessarily included in the model. Depending on the break dates detected by Zivots-Andrew (1992) test, we impose the shift dummy variable ($D_{crisis,t}$) and the impulse dummy variable ($\Delta D_{crisis,t}$), where $\Delta D_{crisis,t} = D_{crisis,t} - D_{crisis,t-1}$. The former captures the shift in the long-run relations, whereas the latter applies for the short-run dynamic models. The VECX* is then given by

$$\Delta y_t = -\Pi_y z_{t-1} + \Lambda \Delta x_t + \sum_{i=1}^{p-1} \Psi_i z_{t-i} + c_0 + c_1 t + c_2 D_{crisis,t} + v_t \tag{5}$$

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_{xi} \Delta z_{t-i} + c_{x0} + u_t \tag{6}$$

with the VARX cointegrating model including a trend term:

$$z_t = (R_{Mt}, R_{Ct}, P_{Mt}, P_{Ct}, EX_{RM/YUANt}, t)' \tag{7}$$

There are $r=2$ cointegrating relations among the 5×1 vector of variables z_t in the conditional model (5) contains three endogenous (Malaysia) variables, $y_t = \{P_{Mt}, R_{Mt}, EX_{RM/YUANt}\}$ and marginal model (6) with two weakly exogenous foreign (China) variables, $x_t = \{P_{Ct}, R_{Ct}\}$. $\Pi_y = \alpha_y \beta'$, α_y is an $m_y \times r$ matrix of error correction coefficients and β' is an $m \times r$ matrix of long-run coefficients and Ψ_i and Λ are the short-run parameters, t is time trend, c_0 is the intercept, and p is the order of VECX*. In the marginal model, Γ_{xi} are the short-run parameters, and c_{xo} is the intercept. It is assumed that u_t and v_t are serially uncorrelated and normally distributed. Notice that we need to restrict the trend coefficients in equation (5) in order to avoid the quadratic trends and the cumulative effects of $D_{crisis,t}$ in the level solution (Pesaran *et al*., 2000), as follow:

$$c_1 = \Pi_y d_1, \qquad c_2 = \Pi_y d_2 \tag{8}$$

where c_1 and c_2 are an arbitrary $m_y \times 1$ vector of fixed constants. Note that d_1 and d_2 are unrestricted if Π_y is full rank; in that case $d_1 = \Pi_y^{-1} c_1$ and $d_2 = \Pi_y^{-1} c_2$. However, if Π_y is rank deficient, d_1 and d_2 cannot be fully identified from c_1 and c_2 but can be estimated from the reduced form coefficients. In this case, the reduced form trend coefficients are restricted. Then, assumes that nominal interest rates, exchange rates, and prices behave in a nonstationary manner. For PPP condition in (1) and UIP condition in (2) to have an empirical meaning, economic theory predicts that:

$$(P_{Mt} - P_{Ct} - EX_{Mt}) \sim I(0) \tag{9}$$

and $$(R_{Mt} - R_{Ct}) \sim I(0) \tag{10}$$

To further justify PPP and IRP, these structural long-run relations require the following (over)-identification restrictions on the cointegration matrix β ($\Pi_y = \alpha_y\beta'$) in equation (5).

$$\beta' = \begin{pmatrix} 1 & 0 & -1 & -1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & -1 & 0 & 0 & 0 \end{pmatrix} \tag{11}$$

where $\beta_1(PPP) = (\beta_{11} \quad \beta_{12} \quad \beta_{13} \quad \beta_{14} \quad \beta_{15} \quad \beta_{16} \quad \beta_{17} \quad \beta_{18})'$

$\beta_2(IRP) = (\beta_{21} \quad \beta_{22} \quad \beta_{23} \quad \beta_{24} \quad \beta_{25} \quad \beta_{26} \quad \beta_{27} \quad \beta_{28})'$

2.2 Data Description

Our analyses are based on monthly observations, spanning from 1994: Jan to 2011: June—a period of economic liberalization and trade expansion for both China and Malaysia. The bilateral exchange rates of RM/Yuan are used in the analyses. An increase of RM/Yuan implies ringgit depreciation against Chinese yuan, and vice versa. For interest rates, the Malaysian base lending rates and Chinese prime lending rates are used. As for price variables, the Malaysian and Chinese consumer prices that adjusted for seasonal effects are compiled and used. All data are sourced from DataStream and cross-checked with the International Financial Statistics, IMF.

3. Empirical Discussion

The preliminary examination of the data properties is conducted using the unit root test of Zivots-Andrew (1992). The data are overwhelmingly integrated of I*(1)* where unit roots are rejected at first difference. This test allows for endogenous structural break, and, for most cases (P_M, R_M, EX_M), the break dates fall on the Asian financial crisis (1997/98) and subprime crisis (2008) periods.[5] We thereby impose two dummy variables on the following long run VARX and error correction VECX* models.

3.1 Long-run Relationship and Restriction Tests

Before proceeding to the cointegration test of long-run relationship, we have to determine the lag orders of endogenous and exogenous variable outlined in Eq (7). For this purpose, the

[5] Results of unit root tests are not presented here but are available upon request.

Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC) are applied to the underlying unrestricted VARX model. AIC has selected the lag orders of 1 for both conditional and marginal models ($k_{SBC}=1, 1$), whereas SBC selected a higher and same order lag ($k_{AIC}=3, 2$) for the endogenous and exogenous variables, respectively. According to Garratt *et al*. (2003) and Affandi (2007), underestimating the lag orders is generally more serious than overestimating them. In Table 1, the log-likelihood ratio statistics that adjusted for small samples (Adj LR) does not reject the VARX model of order (3, 2). As such, the subsequent analyses are based on the VARX (3, 2).

[Insert Table 1]

Next, we need to determine the number of cointegrating relations given by r = rank. The cointegration model contains three domestic variables−P_M, R_M, EX_M, and two foreign variables−P_C, R_C. Following Pesaran, *et al.* (2000), the modified Johansen-Juselius (1992) cointegration test is conducted using λ-max and trace statistics for model with weakly exogenous regressors. The test result is reported in Table 1. It appears that both test statistics indicate the presence of two cointegrating relations ($r = 2$) at 5% significant level based on the bootstrapped critical values by 1000 replications. Such consistent result is in line with the theoretical expectation that PPP and IRP may jointly hold. The PPP relation captures the long run equilibrium of domestic (Malaysia) and foreign (China) prices measured in common currency due to bilateral trading. The IRP relation then captures the equilibrium outcome between domestic (Malaysia) and foreign (China) interest rates due to the effect of the arbitrage process between the two in capital market.

In order to produce the long run estimate of the Malaysia-China parities model, we then impose exact-identifying / normalized restrictions ($\beta_{11} = 1$, $\beta_{12} = 0$, $\beta_{21} = 0$, $\beta_{22} = 1$). In Table 2, the exactly identified ML estimates of the two cointegrating vectors and their asymptotic standard errors are presented. For cointegrating vector one (CV1) that corresponds to PPP, exchange rate and foreign price are statistically significant and carry the expected negative sign. It indicates an established long run PPP relation that goods-market arbitrage will tend to move the exchange rate (RM/yuan) to equalize prices in the two countries. As for CV2 that correspond to IRP, foreign rates of interest also significant and signed correctly, suggesting a potential UIP relationship. UIP states that the financial market (or, the capital account between two currency areas) will only be in equilibrium if, after adjusting for differential risks investors receiving the same rate of return (interest) in both markets. So, if the return on a Malaysia *n*-period interest is one percentage point higher than that on China rate, one would expect, on average, the yuan to appreciate by one percent over the next *n* periods. In addition, exchange rate and foreign prices also plays a significant role in the IRP relation. As for dummy variables, possible positive crisis effect is reported for PPP and negative crisis effect is reported for IRP. Both parity relations are positively affected by the fixed exchange rate regime.

[Insert Table 2]

To further justify the PPP and IRP theorem, we proceed to re-estimate the cointegration relations with seven additional hypotheses using over-identifying restrictions, in addition to the exact-identifying restrictions ($\beta_{11} = 1$, $\beta_{12} = 0$, $\beta_{21} = 0$, $\beta_{22} = 1$). Since LR tests (χ^2) could over-reject in small samples (Affandi, 2007; Garratt, *et al.,* 2006), the bootstrapped critical values based on 1,000 replications of the LR statistic are computed (see Table 3). Using the observed initial values of each variable, the estimated model, and a set of random innovations, an artificial data set is generated for each of the 1,000 replications under the assumption that the estimated version of the model is the true data-generating process.

First, we test the co-trending hypothesis– if the trend coefficients are zero in the two cointegrating relations ($\beta_{16} = 0$, $\beta_{26} = 0$). The bootstrapped critical values for the joint test are 10.49 (95%) and 7.39 (90%) respectively, while the LR statistic (χ^2) of over-identifying restriction is reported as $\chi^2 = 12.01$ in Table 3, hypothesis (a). Hence, the restriction is rejected and the co-trending assumption does not hold. We proceed with the co-breaking hypothesis and the restriction also been rejected, suggesting that PPP and IRP relations are neither co-trending nor co-breaking in the long run. However, in the case of co-pegging, additional restrictions of $\beta_{18} = 0$, $\beta_{28} = 0$ cannot be rejected at 95% and 90% confidence levels. This would imply that the currency pegging to the USD during 1998-2005 do provide supportive evidence for the long run relationships of PPP and IRP between China-Malaysia.

[Insert Table 3]

Next, Eq. 9 suggests that exchange rate (EX_M), foreign price (P_C) and foreign interest (R_C) enter the long run PPP relations with ($\beta_{11} = 1$, $\beta_{13}= -1$, $\beta_{14}= -1$). The reported χ^2 (9.92) is well below the bootstrapped critical values of 12.97 (95%) and 10.24 (90%). Hence, long run PPP holds. Nevertheless, IRP alone does hold when the absolute IRP restriction is imposed ($\beta_{22} = 1$, $\beta_{25}=-1$). The favorable result is observed in (f) when both PPP and IRP are jointly restricted. More important, result in (g) also supports for the cointegrating relationships when the hypothesis incorporated the joint-PPP-IRP and co-pegging restrictions. Overall, our empirical finding confirms the long run validity of joint PPP-IRP for Malaysia-China in the liberalization era. The empirical supports are obtained under the combined assumption that the cointegrating relations are co-pegging but not co-trending or co-breaking. Such finding is in line with Johansen and Juselius (1992), Juselius (1995) and Juselius and MacDonald (2004) that possible interactions between the goods and the capital markets should be allowed to establish the international parity relations.

3.2 *Short run Dynamics and Error Correction Modeling*

Next, what follows is the modeling of VECX* short run dynamics, which is presented in Table 4. Several points are noteworthy. First of all, the lagged error correction terms ($ECT1_{t-1}$ and $ECT2_{t-1}$) for both Price (ΔP_M) and Interest (ΔR_M) equations carry the expected negative and significant sign, indicating that the system - once being shocked, will necessarily adjust back to the long run equilibrium. These estimates shows that the error-correcting coefficient of IRP adjustment is of greater pace in the interest equation ($ECT2_{t-2} = -0.3936$) but slower in the price equation ($ECT2_{t-2} = -0.0034$). On the contrary, PPP adjustment ($ECT1_{t-1} = -0.0578$) is relatively greater than IRP adjustment in the price equation. For price equation, most variables are insignificant, except ΔP_{Mt-1}, ΔP_{Ct-2} and ΔR_{Ct-1}. Then, for interest equation, the lagged ΔP_{Mt-1}, ΔR_{Mt-1} and ΔR_{Mt-2} are significant in explaining Malaysian interest changes. Though with correct signs, the ΔR_{Ct-1} is insignificant in both equations, suggesting rooms of Malaysian monetary autonomy in the short run. Together, the results suggest a direct price transmission from China to Malaysia in the short-run, and Malaysian monetary policy responded to Chinese price to ease domestic inflation. On the other hand, exchange rate does not seem to significantly affect the price changes and interest movements in short-run.

[Insert Table 4]

Despite the R^2 reported as 0.3431 and 0.2946 for the respective price and interest equation in Table 3, three additional diagnostic tests are also conducted. For serial correlation, we use the Lagrange Multiplier (LM) test. The error correction model is clean of autocorrelation problems as the null hypothesis of serial correlation in residuals failed to be rejected, in the presence of lagged dependent variable. The insignificant *F*-statistics are reported at 0.7929 (p-value=0.657) for price equation, and at 1.6171 (p-value=0.100) for interest equation. Using the square of the fitted values, the Ramsey Regression Equation Specification Error Test (RESET) then examines the functional misspecification. The price equation and interest equation are both considered as correctly specified with the *F*-statistics reported as insignificant (p-values=0.121 and 0.756). Likewise, the heteroscedasticity test statistics are also within the insignificant bound for price equation.

3.3 Speed of Convergence and Shock Responses

To this end, it is still imcomplete to conclude how the price and monetary transmission mechanism worked. One should consider the Persistence Profile analysis and generalized Variance Decompositions. In addition to error correction modeling, a good way of measuring the speed of convergence of the cointegrating relations to equilibrium is to examine the dynamic responses of the endogenous variables to various types of shocks. This paper focuses on the effect of system-wide shocks on the cointegrating relations using the Persistence Profile analysis developed by Pesaran and Shin (1996). On impact, the Persistence Profile is normalized to take the value of unity, but the rate at which it tends toward zero provides information on the speed with which the equilibrium correction takes place in response to shocks. In addition to the point estimates, the 2.5% and 97.5% Confidence Bounds—which are generated by employing the nonparametric bootstrap method using 1,000 replications—are also illustrated as dotted lines in Figure 1.

[Insert Figure 1]

The system-wide shock has affected all long-run relations significantly in the beginning, before the effects eventually disappear in the long run. The half-life for PPP relation is about 3.5 months and the whole effect takes around 12 months to complete. The speed of convergence is generally faster than what was documented by Rogoff (1996) but in line with the recent Asian PPP studies (e.g. Baharumshah, Aggarwal and Chan, 2007; Baharumshah, Chan and Fountas, 2008; Chan, Chong and Hooy, 2011). As for IRP relation, the half-life is shown at about 5-6 months and the adjustments completed within a year. The result seems to be consistent with the error correction representation of VECX* model that the convergence process (half-life) in the goods market (PPP) is faster than in financial market (UIP). The faster pace of adjustment (following system-wide shocks) towards price instead of interest equilibrium is also in line with theoretical prediction. Such finding implies the nonappearance of sequencing problem in market integration for Malaysia-China.

Subsequent analysis of the generalized Variance Decompositions (VDCs) is attempting to gauge the extent of shocks to a variable that can be explained by other variables considered in the VARX model. VDCs can be considered as an out-sample causality test, which provides a quantitative measurement of how much the movement in one variable can be explained by other variables in the VAR system in terms of the percentage of forecast error variance. However, the results based on conventional orthogonalized VDCs are found to be sensitive to the number of lag lengths used and the ordering of the variables in the equation. The errors in any equation in a VAR are normally serially uncorrelated by construction, but there may be contemporaneous correlations across errors of different equations. To overcome this problem, we estimate the generalized VDCs of forecast errors (see Pesaran and Pesaran, 1997).

[Insert Table 5]

Table 5 presents the generalized VDCs for our VARX model. Among the five variables in the system, the Chinese variables (P_C and R_C) seem to be the most exogenous variables, as most of the shocks are explained by their own innovations (94%–97% and 92%-99%) over the horizon of 32 months. Such a finding provides the rationale and methodological support to employ the VARX and VECX* modeling in this study. On the other hand, Malaysian price (P_M), exchange rate ($EX_{RM/yuan}$) and interest rate (R_M) are found to be endogenously determined. Nonetheless, the time-lag effect has been evident. The endogeneity of these variables increases by larger proportions after 8^{th} month horizon.

In line with the long-run estimates, innovations from the RM/yuan exchange (>12%), domestic interest (>6%), Chinese price (>5%) and Chinese interest (>4%) explain some portions of the forecast error variance in the Malaysian price (P_M), especially after the 8^{th} month horizon. Apart from the direct effect of imported inflation, exchange rate also plays a significant role in the price transmission mechanism. As for R_M, the major innovation comes from the Chinese price (20%-37%) and domestic price (13%-22%) at increasing rate. Meaning that Malaysian remains the relative monetary autonomy against China but the price channel will affect the extent of IRP condition between the two nations. In addition, about 60%-70% errors in RM/yuan exchange are jointly explained by domestic (P_M), Chinese prices (P_C) and Chinese interest rate. Such finding indicates that the PPP relation is mainly driven by both the price ratio and monetary effect.

4. Conclusion and Policy Implication

Inspired by the work of Juselius (1995), Pesaran *et al*. (2000), and Juselius and MacDonald (2004), this study constructs a structural VARX modeling system that jointly assess PPP and IRP for Malaysia-China, which concurrently allowing *I*(1) exogenous variables in the analysis. Few important findings emerged from our analysis. First, we find overwhelming evidence of both PPP and UIP in the liberalization era (1994-2011), when exchange rate regime and structural breaks were taken into accounts. Second, deviations are shorter lived for PPP. The faster pace of adjustment towards price instead of the interest rate equilibrium implies that sequencing problem in market integration is not an issue. Such supportive empirics are established based on a series of advanced econometric procedures and theoretical formulation which consider possible interactions between the goods and the capital markets. In other words, the present economic linkage provides a platform to promote bilateral free trade agreement, and hence enhancing the closer economic collaboration and financial arrangements for sustainable development.

Nonetheless, the PPP and IRP hold when both China and Malaysia de facto pegged to the USD may also entails with unfavorable economic consequences. The PPP relations implies that any short run deviation of the exchange rates (e.g. real currency depreciation) will be adjusted in the price of tradable goods and hence the trade flows, which steadily revert the exchange rates back to the equilibrium level. But if RM/yuan remains stable within a rigid regime, both PPP and IRP hold to imply that the price hikes will transmit as imported inflation while financial risks are contagious across border. A closer monitor of the Chinese prices and monetary changes is thus essential with the promotion of a more flexible exchange rate between the two nations. And, supply chain diversification would reduce the risk of imported inflation and financial turmoil. In such consideration, our model contributes as an early warning system for Malaysia's economic defense against global shocks.

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Table 1: VARX Cointegrating Tests, 1994Jan-2011June

H_0	$H_{\lambda\text{-}max}$	H_{Trace}	λ-Max statistics	Trace statistics	Bootstrapped Critical Values			
					$95\%_{\lambda\text{-}max}$	$90\%_{\lambda\text{-}max}$	$95\%_{Trace}$	$90\%_{Trace}$
$r = 0$	$r = 1$	$r \geq 1$	86.3183**	130.3805**	38.7363	35.5878	70.7391	65.5733
$r \leq 1$	$r = 2$	$r \geq 2$	34.2299**	44.0621**	30.4143	27.7671	43.4382	40.3196
$r \leq 2$	$r = 3$	$r = 3$	9.8323	9.8323	22.8067	19.7443	22.8067	19.7443
lag(3, 2)				*SBC=1728.1*		*Adj LR test = 119.8041[0.344]*		

Notes: ** and * denote significant at 95% and 90% confidence level respectively. λ-Max statistics are cointegration LR test based on maximal eigenvalue of the stochastic matrix, whereas Trace statistics are cointegration LR tests based on trace of the stochastic matrix. The 95% and 90% critical values are generated by bootstrap method using 210 observations and 1000 replications. The underlying International Parity VARX model contains unrestricted intercept with trend and the optimal lag order (3, 2) based on SBC is shown at the bottom of Table 1.

Table 2: Exact-identifying Restrictions, 1994Jan-2011June

Exact-identifying Restrictions ($\beta_{11} = 1, \beta_{12} = 0, \beta_{21} = 0, \beta_{22} = 1$)								
	P_M	R_M	$EX_{RM/yuan}$	P_C	R_C	T	D_{98}	D_{FIX}
CV1(PPP)	1.000	0.000	-0.2838** (0.0533)	-0.1759** (0.0284)	-0.0745** (0.0284)	-0.0013** (0.0001)	0.0506** (0.0123)	0.0107** (0.0047)
CV2(IRP)	0.000	1.000	0.9144* (0.4681)	1.3911** (0.4423)	-0.3461* (0.1947)	0.0029** (0.0012)	-0.3369** (0.1093)	0.1049** (0.0264)

Notes: ** and * denote significant at 95% and 90% confidence level respectively. CV1 and CV2 represent the respective cointegrating vector for PPP and IRP. Asymptotic standard errors are reported in the parentheses. Dummies for subprime crisis (D_{08}) were found insignificant and omitted from both models.

Table 3: PPP and IRP Restriction Tests, 1994Jan-2011June

Hypotheses		Exact-identifying + Over-identifying Restrictions	LR (χ^2)	Bootstrapped Critical Values	
				95%	90%
(a)	co-trending	$\beta_{16} = 0, \beta_{26} = 0$	12.0147**	10.4928	7.3964
(b)	co-breaking 98	$\beta_{17} = 0, \beta_{27} = 0$	39.6945**	9.4023	7.2139
(c)	co-pegging	$\beta_{18} = 0, \beta_{28} = 0$	5.7822	7.7878	6.1559
(d)	PPP	$\beta_{13}= -1, \beta_{14}= -1, \beta_{15}=0$	9.9218	12.9705	10.2456
(e)	IRP	$\beta_{23}=0, \beta_{24}=0, \beta_{25}=-1$	13.5658**	11.5540	10.3824
(f)	PPP+IRP	$\beta_{13}= -1, \beta_{14}= -1, \beta_{15}=0, \beta_{23}=0, \beta_{24}=0, \beta_{25}= -1$	15.1527	19.3559	15.8670
(g)	PPP+IRP+ (c)	$\beta_{13}= -1, \beta_{14}= -1, \beta_{15}=0, \beta_{18}=0, \beta_{23}=0, \beta_{24}=0, \beta_{25}= -1, \beta_{28}=0$	17.3866	23.4172	20.4195

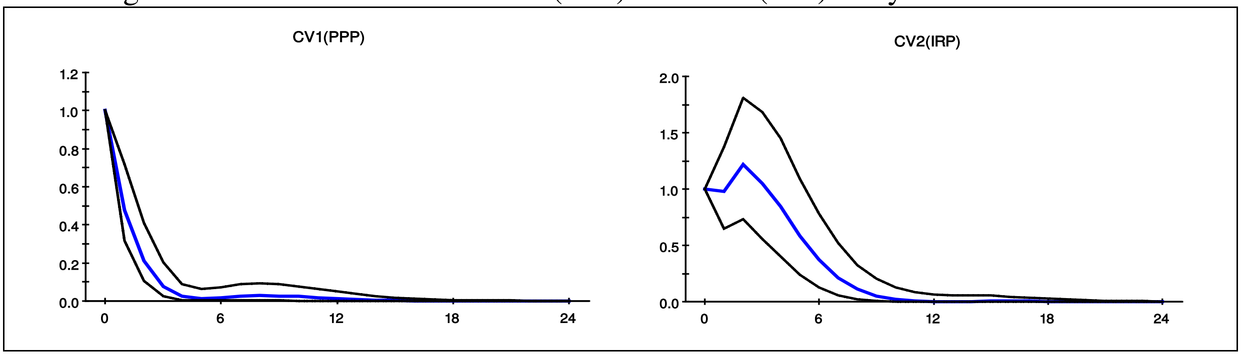
Notes: ** denotes significant at 95% confidence level. The respective 95% and 90% critical values are generated by bootstrap method using 210 observations and 1000 simulations. All ML estimates converged within 100 iterations. The underlying VARX trade model is of lag order (3, 2) and contains unrestricted intercept with trend.

Table 4: Error Correction Representation in VECX* Modeling

Regressor	Dependent Variable ΔP_M	ΔR_M
	Conditional Model	
ΔP_{Mt-1}	0.2143 (0.0697)***	0.3791 (0.4523)
ΔP_{Mt-2}	0.0920 (0.0695)	1.1755 (0.4510)**
ΔR_{Mt-1}	0.0088 (0.0101)	0.2486 (0.0655)***
ΔR_{Mt-2}	-0.0109 (0.0101)	0.1492 (0.0654)**
ΔEX_{Mt-1}	-0.0065 (0.0073)	-0.0671 (0.0476)
ΔEX_{Mt-2}	0.0084 (0.0074)	0.0360 (0.0482)
c	0.1996 (0.0745)	1.0912 (0.4831)**
$ECT1_{t-1}$	-0.0578 (0.0218)***	-0.0849 (0.0176)***
$ECT2_{t-1}$	-0.0034 (0.0027)	-0.3936 (0.1415)***
	Marginal Model	
ΔP_{Ct-1}	0.0317 (0.0548)	0.5708 (0.3556)
ΔP_{Ct-2}	-0.7970 (0.3586)**	-0.0056 (0.0553)
ΔR_{Ct-1}	0.0210 (0.0095)**	-0.0526 (0.0613)
ΔR_{Ct-2}	-0.0020 (0.0094)	-0.0278 (0.0609)
	Diagnostic Tests	
R^2	0.3431	0.2946
AUTO	0.7929[0.657]	1.6171[0.100]
RESET	2.4260[0.121]	0.0966[0.756]
Hetero	0.7750[0.380]	4.0596[0.045]

Notes: *, **, *** denote significant at the 10%, 5%, and 1% levels, respectively. AUTO is the Lagrange Multiplier test for serial correlation; RESET is the Ramsey Regression Equation Specification Error Test for functional form; and Hetero tests for heteroscedasticity. All diagnostic tests are conducted using *F*-statistics. Standard errors and p-values are presented in ( ) and [ ] respectively.

Figure 1: Persistent Profile of CV1 (PPP) and CV2 (IRP) to System-Wide Shocks



Note: The dot-lines represent the top 97.5% and low 2.5% bootstrapped confidence intervals respectively.

Table 5: Generalized Variance Decomposition for VECX* Model

Variables	Horizon	% of Forecasted Variance Explained by Innovations in				
		P_M	R_M	$EX_{RM/yuan}$	P_C	R_C
P_M	1	**92.44**	0.03	4.12	0.51	2.89
	4	**85.01**	0.30	10.18	0.90	3.61
	8	**80.31**	1.38	12.62	1.63	4.06
	16	**74.08**	4.88	12.56	3.88	4.60
	24	**70.81**	6.86	12.05	5.35	4.92
	32	**69.20**	7.80	11.82	6.09	5.09
R_M	1	0.02	**98.65**	0.11	0.74	0.48
	4	0.43	**97.49**	0.79	0.61	0.69
	8	1.56	**92.93**	0.97	3.39	1.14
	16	13.10	**62.90**	2.07	20.57	1.37
	24	19.74	**44.05**	3.40	31.60	1.21
	32	22.87	**34.97**	3.97	37.06	1.14
$EX_{RM/yuan}$	1	3.84	0.20	**87.97**	6.15	1.84
	4	14.60	0.79	**65.25**	12.78	6.58
	8	25.05	4.02	**48.79**	12.20	9.93
	16	34.60	9.38	**34.14**	9.59	12.30
	24	40.97	9.69	**27.40**	8.85	13.10
	32	44.67	9.64	**23.59**	8.56	13.53
P_C	1	0.04	0.02	0.88	**97.87**	1.19
	4	0.20	0.08	0.90	**95.81**	3.01
	8	0.40	0.05	0.60	**95.05**	3.91
	16	0.46	0.08	0.34	**94.67**	4.44
	24	0.48	0.09	0.25	**94.57**	4.60
	32	0.50	0.09	0.21	**94.53**	4.67
R_C	1	0.07	0.00	0.01	1.00	**98.92**
	4	0.22	0.00	0.10	3.78	**95.90**
	8	0.30	0.00	0.12	5.52	**94.06**
	16	0.32	0.03	0.10	6.40	**93.16**
	24	0.32	0.04	0.08	6.64	**92.91**
	32	0.33	0.04	0.08	6.77	**92.78**