# The Long-run Dynamics of Interaction between China's Onshore and Offshore Exchange Rates

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2016

*Abstract*: This paper addresses a key question of relevance to the Chinese exchange rate unification, namely whether its current dual onshore and offshore rates are co-integrated. The paper explores the long and short-term dynamics of the relationship between the time-series of onshore and offshore exchange rates through a new approach based on the differential between the two rates. Based on the evidence of cointegration of the two rates, we test the null hypothesis of a linear two-equation vector error correction (VEC) model against nonlinearlity, and further estimate a two-regime Markov-switching chain model to evaluate the predictive power of the exchange rate differential for the behavior of China's dual exchange rate behavior. We find that the one-period lagged differential displays statistically significant short-term disequilibria correction effects with respect to the long-run equilibrium path of the differential rate. The Lagrange Multiplier (LM) tests for ARCH effects on the regression residual of the black market exchange rate on the official exchange rate suggest volatility; we also report estimates by ARCH processes for ARCH, and for GRACH effects. The study suggests the exchange rate differential approach can provide an effective predictive and forecasting tool for China's full unification policy assessment.

## I. Introduction

The literature on exchange rate analysis in parallel markets has relied extensively on the autoregressive effects of the black market and official exchange rates in developing models and forecasts for exchange rate time-series. For example, in a large scale study of black market exchange rates of 32 economies, Gramacy *et. al.* (2014) found that the official exchange rate is the second most important "fundamental" determinant of the black market exchange rate. Surprisingly, however, has been the relative absence of the exchange rate differential in time-series analysis and forecasting in the literature. One would expect the variation of the lagged exchange rate differential to reflect the influences of all principle fundamentals that affect the supply and demand imbalances of the exchange rate in markets with a parallel rate structure; therefore, it may be a promising tool of analysis in such markets. This is not to deny the important but . limited role that the concept of the black market premium . has played in the literature. The Dornbusch et. al. (1983) seminal paper has influenced the exchange rate parallel market literature greatly but that approach models a *partial* equilibrium demand for currency; a more fully articulated *macro-market* analysis appears to be missing, see, however, Phylaktis and Kassimatis (1994).

This paper is addressed to this gap, demonstrating the incorporation of the differential rate approach into standard econometric time-series models. We conduct our analysis on the basis of the cointegration of the black market and official exchange rates, and use that outcome to apply vector error correction (VEC) models based on the differential rate, and develop a framework for modeling the interaction of the short-term and long-term exchange rate dynamics. We further examine the volatility of the residual variance obtained from the exchange rate differential with estimation obtained from ARCH dynamic processes.

We apply that approach to a long weekly time-series of the differential rate between the onshore and the offshore Chinese exchange rates covering 8/23/2010 to8/12/2016. Our results support the existence of a long-run cointegrated relationship between the two exchange rates, and our estimates of a linear and two-regime Markov VEC models indicate a significant unit-time predictive role of the differential variable on the exchange rate time-series. Moreover, our LM tests reveal the nonstationarity of the second moment of the differential residual and the presence of significant ARCH effects. This is a preliminary and partial study of Iran's exchange rate time-series based on the differential rate approach intended to be developed further to compare the forecasting error of the differential-rate-based approach with fundamental-based approaches.

Section II reviews .the exchange rate time-series literature that bears, directly or indirectly, on the parallel currency markets; section III sets out the statistical tests and econometric model specifications relevant to a differential rate approach, and briefly discusses the data sets in section IV. Section V presents the outcome for implementing the econometric approach outlined in section 3, and a final section VI presents the conclusion.

## II. Literature Review

In a seminar paper, Dornbusch *et. al.* (1983) provided a model for black market exchange rate behavior that influenced much of the research in subsequent periods. For example, Fishelson (1987) employed that model to demonstrate that the expected speculative profit and the official exchange rate act as the key determinants of the black market exchange rate. However, the forecasting ability of the Dornbusch model remained unexplored in the literature; more important, its analysis is focused on the partial equilibrium demand for foreign currency. In an equally well-known paper, Meese and Rogoff (1983) examined the forecasting ability of structural models of the exchange rate and found them no better than a simple random walk model; subsequent nonlinear structural models reinforce that outcome. Much of these later developments follow the approach of analyzing the black market exchange rate as a function of fundamentals such as GDP, inflation or money supply, see Engel and West (1994). Engel (2004) applied a Markov chain model, see Hamilton (1991) for an exposition, to the exchange rate as a function of economic fundamentals but failed to obtain superior exchange rate forecast values for the model, although they found weak support in the model's directional forecast ability. A number of studies, however, have exploited time-series for many countries combined into a larger panel data set to overcome the weaknesses of the structural models suggested by Meese and Rogoff. Mark and Sul (2001) reported panel co-integrated relationships between the exchange rate and fundamentals, and superior forecasting accuracy compared with a random walk, based on . panel time-series for a sample of developed economie. Cerra and Saxena (2008) also employed a similar approach using a much large data set of developed and developing countries with periods of parallel exchange rates containing a much lower degree of cross-sectional correlation; yet reported similar results. Gramacy *et. al.* (2014) analyzed the structural models of black market exchange rates for a large sample of countries employing nonparametric, nonlinear, Markov chain methods, and demonstrated that such models have much better forecasting ability compared to random walk models. They found speculative activity and, to a lesser extent, the official exchange rate. are the two most important determinants of the black market exchange rate across a large number of countries. Hence, the fundamentals-based approach has been a dominant feature of the literature; for example, both Mark and Sul, and the Cerra and Saxena test for cointegration of the exchange rate with the als, and employed forecasting models with deviations of the exchange rate from the fundamentals, . their lagged values, as independent variables.

This study examines a relatively novel alternative approach to the determinant of the exchange rate in parallel currency markets based on the information contained in the exchange rate series themselves. As Information about the fundamentals with significant influence on the black market exchange rates behavior is likely to affect the difference between the black market and official rates, that gap should contain at least as much information as key fundamentals do. A test of the hypothesis that the long-run equilibrium exchange rate gap tends .toward zero, provides a test for the cointegration of the two rates, see Engle and Granger (1987). Provided cointegration is confirmed, a one-period deviation of the black market exchange rate from the official rate, that is, the lagged error term of the regression of the black market rate on the official rate, represents short term divergence from equilibrium. The error correction model presents an econometric model for long-run dynamics with a short-term disequilibrium error correction term embedded within the model as the key determinant of adjustment in the direction of equilibrium. We employ this framework to analyze a dual exchange rate market. A statistically significant coefficient estimate for this variable provides primary evidence of the predictive power of the exchange rate deviation in explaining the exchange rate behavior in a parallel market[1]. The absence of an exchange-rate-gap approach in the parallel market literature is surprising, given its potential predictive power; however, Phylaktis and Kassimatis (1994)

[1] Indeed, such an approach is even more promising in other parallel markets with non-sticky open-market prices if the cost of price adjustment is negligible; therefore, adjustment to equilibrium is via price, rather quantity change, notably in food markets, see Koohi-Kamali (2012).

appear to be an exception. They have examined a two-equation, linear vector error correction model with the exchange rate gap as the error correction term. However, a typical long time-series of official exchange rate is likely to consist of several periods of sharply different exchange rate regimes; making a non-linear modelling of the exchange gap essential. In this paper we go beyond a linear VEC model to examine the predictive power of the exchange gap with a Markov chain VEC model.

Cointegration tests are focused on examining the stability of the first moment in a time series. Exchange control policies are often adopted with the aim of moderating the volatility of the exchange rate; hence examination of the second moment of the time-series of exchange rates has been another focus of the literature on exchange rates in parallel markets. A volatile time series has a leptokurtic distribution; with a fatter tail than the normal. One question is whether the shape of such distributions is affected by the time-frequency of the data. Diebold (1988) demonstrates that under temporal aggregation (from daily or weekly to monthly), the ARCH processes converge.to the normal distribution. It is thus an important issue whether the application of ARCH models to monthly time-series can reveal evidence of volatility that may otherwise be clear with daily or weekly data sets. Both these questions are addressed in Phylaktis and Kassimatis (1997) who apply ARCH and GARCH models to monthly time-series of four Pacific Basin economies by also adding a dummy variable that controls for sharp policy changes from period to period. This study fails to reject the unit-root hypothesis for both series, and thus applies the ARCH model to the first differences of each rate, and found relaxation of the official rate .tends to increase exchange rate volatility, while controlling for policy change tends to reduce it. Therefore, the evidence of volatility in terms of high kurtosis is clear in this study even though the data sets employed consist of monthly data.. The major shortcoming of this study is the application of ARCH models to the official and black market rates separately. If, as the evidence in this study suggests, the gap between the rates tends to be a cointegrated series, then there would be no need to apply the ARCH models to first-differenced rates. The ARCH processes should be applied directly to the variance of the error term obtained from the regression of the black market rate on the official rate; that is, to the gap between the two rates. This study, however, is somewhat different to the ARCH applications discussed above; and in any case, fails to reject the hypothesis that the variance of the differences between the rates.  is stable. The neglect of this issue is all the more puzzling,, as the evidence of the ability of the ARCH effect terms to forecast volatility with reasonable accuracy should be of major interest.

China, whose exchange rate scheme is examined in terms of the above approach, has gone through a number of different controlled official exchange rates since the late 1970's. In 1986 swap exchange rate centers were established that allowed Chinese businesses to trade with foreign investors at pre-determined rates, which differed from the official rates. By 1994, the official and swap rates were unified at the prevailing swap rate. Phylaktis and Girardin (2001) take a fundamental. approach to the examination of the relationship between the Chinese black market and swap rates using quarterly data sets covering 1988to 1996 and estimate a regression of the black market rate on the official rate, the ratios of the GDP and of inflation inChina, compared  to the  USA, all expressed in logarithmic terms. Their tests of stationarity confirm that the fundamentals are co-integrated. Note that the EC term estimates for the two exchange rate equation contain the exchange rate gap as a component, in addition to information on GDP and inflation.  Therefore, it is not clear from this study how much of the 42 % correction toward the equilibrium value of the black market rate can be attributed to the dual exchange rate gap. Their impulse-response analysis shows the swap rate derives much of the movement in the other variables. The main conclusion of this study is that the output affects the black market rate through the demand for money; and that the swap rate is exogenous in the black market equation; and output has a large negative impact on the black market rate. The official implementation of an offshore, open market exchange rate, initially working through the Hong Kong market, in August 2010 is the focus of the Li et al. (2012) analysis of the interaction between the onshore (CNY) and offshore (CNH) exchange rates. This study explores the additional causes of the offshore and onshore differential beyond that explained by the separation of the two rates through official control, or management, of the onshore rate, using monthly data over the period of 2006-2011. They test the role of uncertainty created by the difference in the available information for investors in each market, due to the limits imposed on trade for the offshore investors. The authors find that after controlling for the effects of the fundamentals (lagged exchange rate differential, market volatility risk, capital flight, etc.) on the offshore/onshore exchange rates gap, a 1 percent increase in parameter uncertainty leads to 54.3 basis point change in that gap. Cheung and Rime (2014) examine the effect of directional long and short betting, with order flow data available from microstructure daily data sets, to examine the onshore/offshore rates interaction.

The tests confirms cointegration of the two exchange rate series; the authors employ an augmented VEC to model long and short term dynamics of China's exchange rate and obtained a significant EC term for the offshore equation equal to 0.65539 ($t$=4.21), i.e. a 65.6 % correction for the gap between the offshore rate and its long-run mean value in every unit of time EC term is insignificant in the onshore equation). This study finds that the order flow variable explains respectively, 29% and 39% of the change in the onshore and offshore series respectively. However, the extent of the effect of the rate gap on each series remains unknown since the EC estimates contain that effect only as a component. The authors also provide out-of-sample, one-period-ahead forecasts for the official daily central parity renminbi rate obtained the regression of the latter on the first-differences of the logs of CNH, CNY, and the order flow variable and test the root mean squared error and absolute squared error for the forecast of each regression against those obtained from a random walk model with a drift. They demonstrate that the model with the order flow as the dependent variable .provides the forecasts with the smallest error. It is clear that there are also other influences that affect onshore and offshore exchange rate markets' behavior. Maziad and Kang (2012) application of the ARCH processes to Chinese onshore and offshore exchange rate volatility do allow for the spill-over effects between the two rates, but that study is not an application of ARCH processes to the time-series of the exchange rate differential. Liu (2015) employs high-frequency Chinese onshore-offshore time-series data and reports the presence of significant asymmetric bad and good news effects on the differential rate. This study, however, is somewhat different from the ARCH applications discussed above; and in any case fails to reject the hypothesis that the variance of the differential rate is stable.

## III. Econometric Model Specification

In this section we outline VEC model specifications that are tied to the basic characteristics of China's dual exchange rate. Consider the following ARIMA model for the basic relationship between the offshore exchange rate (*HR*) and the onshore exchange rate (*YR*) from which the EC model can be derived by imposing the parameter restrictions implied by that model.

$$ln\_HR_t = \beta_0 + \beta_1 ln\_YR_t + \beta_2 ln\_YR_{t-1} + \beta_3 ln\_HR_{t-1} + e_t \qquad (1)$$

Solving for long-term equilibrium ($ln\_HR_t = ln\_HR_{t-1}$; $ln\_YR_t = ln\_YR_{t-1}$) lead to the EC model

$$\Delta ln\_HR_t = \beta_{1\Delta} ln\_YR_{t-1} + (\beta_3 - 1)\, e^*_{t-1} + e_t \qquad (2)$$

$e^*_{t-1}$ in (2) represents the one-period EC disequilibrium term. The relevance of a drift term is decided empirically with reference to the graph of the differential rate given in plot 1 that displays fluctuations around a horizontal line drawn at zero mean. This suggests the *exclusion* of a drift term for the specification of $e^*_{t-1}$ equation[2].

$$e^*_{t-1} = (ln\_BR_{t-1} - \theta\, ln\_OR_{t-1})\ ;\ \theta = -(\frac{\beta_1 + \beta_2}{\beta_3 - 1}) \qquad (3)$$

It can be shown that the hypothesis for the two rates having the same rate of change implies $\theta = 1$;that is, by imposing the parameter restriction ($\beta_1 + \beta_2 + \beta_3$)=1 on eq. (1), see Hendry *et. al.* (1992, pp.140-8). The single-equation EC model presupposes exogeneity of the RH variables; hence we employ the following two-equation VEC model augmented with the lagged values of variables in first differences, and simplified by including only the first-order lags suggested by our empirical results, see below.

$$\Delta ln\_HR_t = \beta_{01} + \beta_{11} ln\_YR_{t-1} + \Delta\beta_{21} ln\_HR_{t-1} + \Delta\beta_{31} ln\_YR_{t-1} + \beta_{41} ln\_HR_{t-1} + \theta_1 e^*_{t-1} + e_{t1} \qquad (4)$$

$$\Delta ln\_YR_t = \beta_{02} + \beta_{12} ln\_YR_{t-1} + \Delta\beta_{22} ln\_HR_{t-1} + \Delta\beta_{32} ln\_YR_{t-1} + \beta_{42} ln\_HR_{t-1} + \theta_2\, e^*_{t-1} + e_{t2} \qquad (5)$$

The $e^*_{t-1}$ is the EC term common to both equations whose coefficient, $\theta_i$ for $i$= 1, 2, is a measure of the one-period disequilibrium correction toward the mean of the cointegrated series. We expect $\theta_1<0$ in (4) and $\theta_1>0$ in (5), see below.

An important question to address is whether the relationship between the offshore and onshore exchange rates can be adequately modeled by the linear VEC model in (4) and (5). Such an approach faces a parameter identification problem; that is, when each regime is defined by a

[2] Phyliktis and Kassimatis (1994) advocate a differential rate exclusive of a drift term arguing that the Dornbusch *et. al.* (1983) stock-flow model that implies an equilibrium black market premium rate defined by the black market exchange rate *spread* excluding a drift term. This specification happens to be also supported in plot 1.

different time threshold, the threshold coefficient will be the same in a linear model, and thus not identified. Time-series of parallel market exchange rates over a long period are subject to changing periods of official exchange rates, and therefore, involve different exchange rate regimes as is clear from plot 1. Our tests for nonlinearity are based on the null hypothesis of a linear model against a non-linear alternative nesting the linear specification. The literature, especially on smooth transition models, recommends a Taylor expansion around parameters of the transitional variable, typically the lags of the dependent variable, to test for the additional polynomial and interactive terms that arise over and above the linear specification, see Elliot and Timmermann (2016, p.180), Gonzalez-Rivera (2013, ch. 16), and below for the implementation of this test.

Based on the graph of the differential rate given in plot 1, we further employ a three-state Markov switching chain version of the linear VEC models of (4) and (5) specified as

$$\Delta ln\_HR_t=\beta^{S1t}{}_{01}+\beta^{S1t}{}_{11}ln\_YR_{t-1}+\Delta\beta^{S1t}{}_{21}ln\_HR_{t-1}+\Delta\beta^{S1t}{}_{31}ln\_YR_{t-1}+\beta^{S1t}{}_{41}ln\_HR_{t-1}+\theta_1{}^{S1t}e^*{}_{t-1}+e_{t1} \quad (6)$$

$$\Delta ln\_YR_t=\beta^{S2t}{}_{02}+\beta^{S2t}{}_{12}ln\_YR_{t-1}+\Delta\beta^{S2t}{}_{22}ln\_HR_{t-1}+\Delta\beta^{S2t}{}_{32}ln\_YR_{t-1}+\beta^{S2t}{}_{42}ln\_HR_{t-1}+\theta_2{}^{S2t}e^*{}_{t-1}+e_{t2} \quad (7)$$

where $s_{1t}$ = $ln\_HR_{t-1}$ and $s_{2t}$ = $ln\_YR_{t-1}$ g are the transitional variables. The substitutions of the transitional variables $s_{1t}$ and $s_{2t}$ in (6) and (7) lead to the Markov switching model specifications employed in the estimations reported in section V below, see section V below for its implementation. The first-order Markov regime-switching assumes that the transitional probabilities of regime change depend only on the move from the previous regime, that is, moving from a recession period to one of expansion, . is independent of how long the recession has lasted. They are therefore, constant, for moves from $i$ to $j$ regimes and expressed as expressed as $P_{ij}=P(s_t=i|s_{t-1}=j)$ and $\sum_{j=1}^{3} p_{ij} = 1$ *for all i* . Since the transitional probabilities sum up to one, we need to estimate a subset of the probability matrix. For example, for a two-regime Markov chain matrix, $P=\begin{pmatrix} p_{11} & 1-p_{22} \\ 1-p_{11} & P_{22} \end{pmatrix}$, only $P_{11}$ and $P_{22}$ estimates are required.

So far, the focus has been on stationarity of the mean of the two series; however, mean-stationary series may have non-stationary variance. In this study, we also test stationarity of the second moment of the exchange rate differential. We employ the simplest equation for China selected by Augmented Dicky-Fuller (ADF) tests as the model that has white-noise error.

$$h_t = \delta + \alpha_1 e_{t-1}^2 + \alpha_2 e_{t-2}^2 + \beta_1 h_{t-1} \tag{8}$$

where $e_t$ is the disturbance term from the regression of $\Delta ln\_HR_t$ on $\Delta ln\_YR_t$, without a drift term, and $h_t$ stands for its variance; with necessary conditions for covariance stationary nature, $\delta > 0, \alpha_i \geq 0, \beta_i \geq 0, (\Sigma\alpha_i + \Sigma\beta_i) < 0$.

We furthermore test the structural stability of this model by adding a dummy variable that divides the series depicted in plot 1 into the two exchange rate regimes. The selection of a break-point is based on the lowest point of the onshore series, which occurred on February. 17, 2014. Eq. (8) is first estimated on the residual obtained from the regression of $\Delta ln\_HR_t$ on $\Delta ln\_YR_t$, in this case without a drift term, and then repeat the regression by adding the time-break dummy variable to the equation, and estimate (8) with the new residual series to test for ARCH effects and estimates the size of those effects.

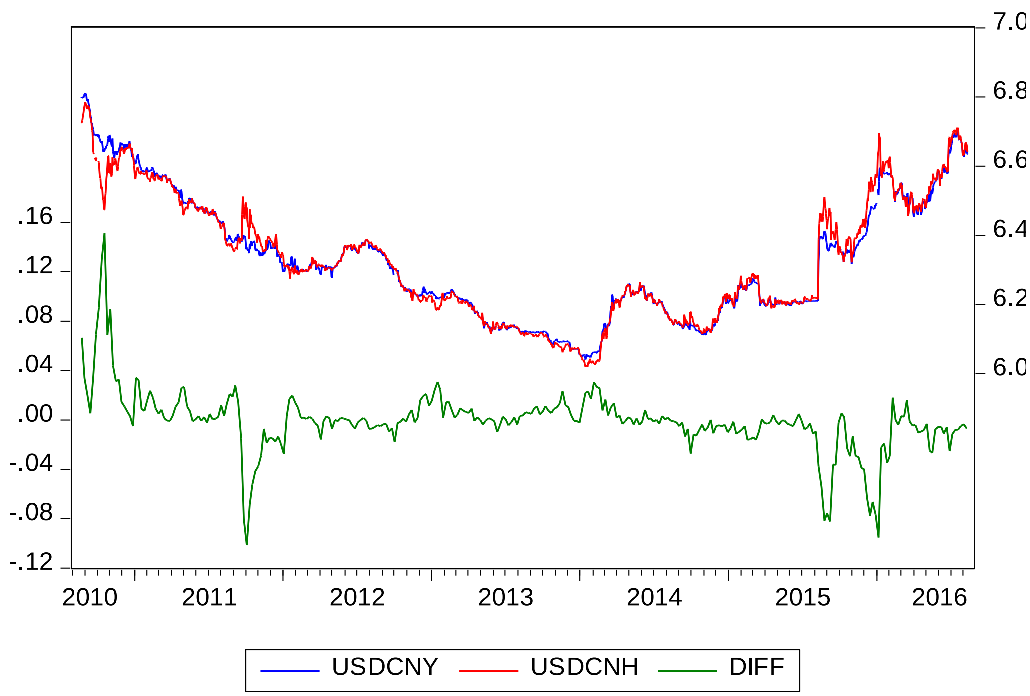
**IV Data**

China has a long history of adopting parallel rates for her exchange rate policies, going back to the late 1970's. However, we focus on a much shorter series starting with the establishment of the open-market exchange in Hong Kong in 2010 since we feel the more recent period is of greater interest to policy debates about China's full exchange rate unification path. This choice is made easier by the availability of exchange rate data at high-frequency that allow aggregation to levels that would still provide for time-series of sufficient duration for the purposes of this study.

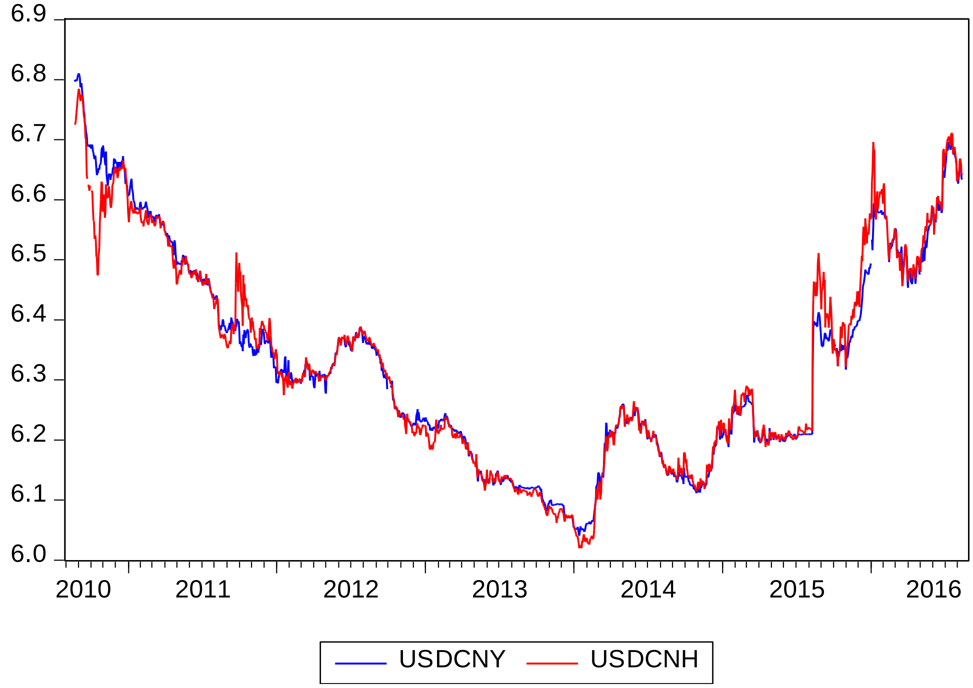
The data employed for this study is from the Bloomberg Terminal, the source that offers. the longest history of offshore rates. USD/CNY, or Dollar-Yuan onshore,; is available from 1/1/1957onward, and can be updated daily. USD/CNH, or Dollar-Yuan offshore (London Exchange was where it received the official designation), is available from 8/23/2010 onward, and can be updated daily. If without the addition of other data, we will take the sample from 8/23/2010 to date (11 days later, 8/22/2016, we will have had a full 6 years of data).

We can further decompose them into very high-frequency intraday series if necessary; or aggregate into weekly, semi-monthly, monthly, quarterly, and yearly terms, using the averaging method. . There are 1558 entries in the daily series, 313 entries in the weekly series, 73 entries in the monthly series; and 6 entries in the yearly series. The weekly series would be a large enough sample size to carry out the study. This is also a good choice in the sense that Chinese authorities tend to have a habit of announcing policy changes on a Monday – the specific weekly entry should capture the full effect and distinguish each period from those of the previous week. Right after a policy adjustment or new policy comments, we often observe the offshore rate over- or undershoot the onshore rate, which is also evident from enlarged graph of both given in plot 2. The onshore adjustment then converges towards the onshore rate in the following period. Plot 1 displays the weekly series for the onshore-offshore series from 8/23/2010 to. 8/12/2016; nearly a full six years of exchange rate time-series.

***Plot* 1** Weekly time-series of China's differential, onshore and offshore exchange rates (in levels) 2010:1 to 2016:1



***Plot* 2** Weekly time-series of China's on-offshore exchange rate (in levels) 2010:1 to 2016:1




## V. Results

We employ the ADF test and the Schwarz criterion selects one lag for each series $ln\_HR_t$ and $ln\_YR_t$ to remove all autocorrelation effects from the disturbance term. The results, shown in table 1, indicate the hypothesis of no unit(?).is rejected at all acceptable levels. Hence both series have nonstationary $I(0)$, which is also evident from insignificant *p*-values.

To test for co-integration, we regress the black-market exchange rate on the official rate series, and the first lags of both series suggested by the Schwartz model selection criterion. The ADF test statistic (-5.11), last row of table 1, is highly significant at the 1% DF critical value; a significant *p*-value now indicates a white-noise disturbance term. Hence, the test result shows the open-market *HR* and the managed *YR* series are co-integrated, that is to say, in the long-run both change in the same proportion. Since by the Granger representation theorem, Engel and Granger (1987), the relationship between the two rates has an equivalent error correction representation; we apply a VEC model to estimate the interactions between the long and short-term dynamics of that relationship.

We first estimate a linear two-equation VEC model, . to test the null hypothesis . against nonlinear VEC processes. The model selection rule picks a VEC model with only the first-order lag-s for each exchange rate, that is, eq.s (4) and (5). As a first approximation, such a linear model is not without interest in providing preliminary evidence of the predictive power of the lagged differential rate, shown in the first column of table 2. The error correction coefficient estimates are statistically significant at the 5% level for $\Delta ln\_HR_t$, at the 10% level for $\Delta ln\_YR_t$, and have the expected sign in both equations: negative in (3), and positive in (4). These suggest that in any single period, the short-tern dynamic lowers $\Delta ln\_HR_t$ by 1 percent and raises $\Delta ln\_YR_t$ by 5.6 percent to bring the two rates closer to their long-term equilibrium path. However, only $\Delta ln\_HR_{t-1}$ in (4) has a significant lagged impact.

We test this linear VEC specification with a single exchange rate regime against nonlinearlity that arises from having changing regimes. A transitional function from one regime to another is defined as a function of the transitional variable; since only the first lags are significant, we take $\Delta ln\_BR_{t-1}$ and $\Delta ln\_OR_{t-1}$ as the transitional variables, and expand the linear VEC around zero of the slope parameters. We expand (4) and (5) by including .the cubic terms for the transitional variables together their interactive terms, that is, we add nine additional terms to (4) and (5) and test if these are collectively all equal to zero. The $\chi^2$ Wald test statistics (35.2) for df=9 and the $F_{(9, 297)}$ test statistic (3.9) both reject the linear VEC specification for (4) at the 5 % level, but for (5), the test-statistics are insignificant and fail to reject linearity, see the second column of table 2, last row. .Therefore, for eq. (4), a linear VEC is a misspecification.

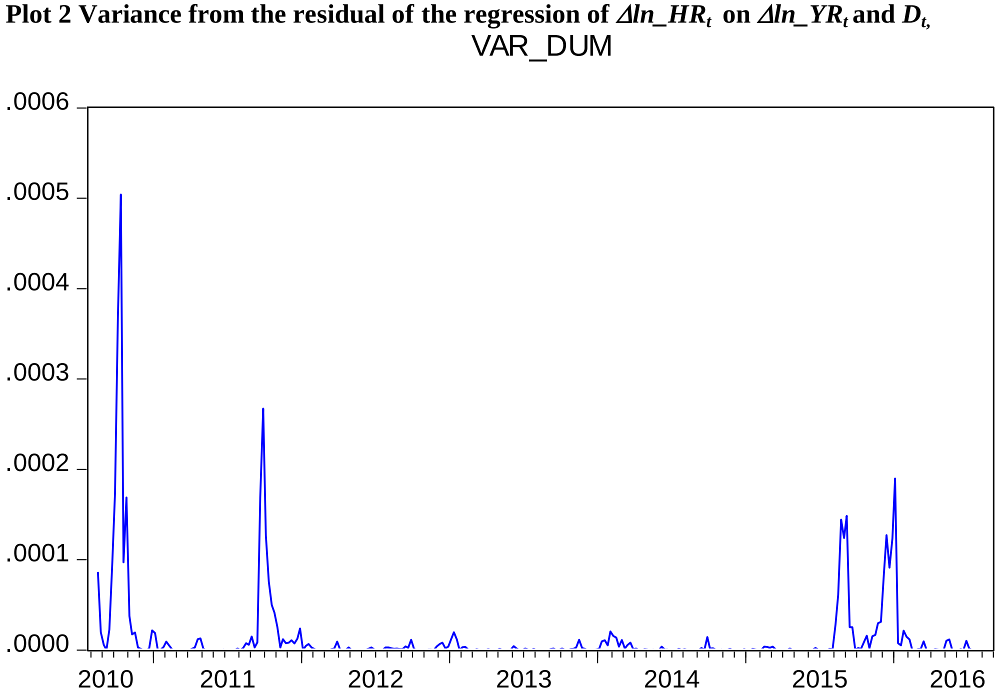
Next, in table 3, we present our nonlinear estimates corresponding to the above two-equation VEC system by a two-regime Markov-switching chain process specified in (6) and (7). We note that in both equations, there is 1% level, statistically significant error correction term in regimes one and two, though somewhat weaker for $\Delta ln\_YR_t$. There are also some leg effects, but the EC term is the main driver of the two rates towards equilibrium. Transitional *p*-values in .regime two for $\Delta ln\_YR_t$ (0.989) indicates a highly persistent regime; that is, there is high probability that $\Delta ln\_YR_t$ in regime two will be followed by another from the same regime. This is reflected in the constant expected duration of $\Delta ln\_HR_t$ and $\Delta ln\_YR_t$, obtained from $1/(1-P_{ii})$ for $i$=1, 2. 3: 1.78 to 1.46 weeks for $\Delta ln\_HR_t$, and from 2.6 to nearly 100 weeks for $\Delta ln\_YR_t$.

Next, given that the disturbance from the regression of $\Delta ln\_HR_t$ and $\Delta ln\_YR_t$, (without a drift) has a stationary mean, we examine the volatility of its second moment. Table 4 reports the Lagrange Multiplier (LM) test for ARCH effects for the first-order autoregessive of the square. of the disturbance term obtained from the regression of $\Delta ln\_HR_t$ on $\Delta ln\_YR_t$, and then again by

An addition of a time-dummy variable to that regression that corresponds to the lowest value of the onshore series in plot 1 (Feb./17/2014). We also estimate (8) with the new residual. The $\chi^2$ LM test statistics, reported in the last row, decisively reject the null hypothesis of no ARCH effects. We estimated several families of ARCH models for the residual variance of the regressions employed in table 4, with different lag structures, including an ARCH threshold model with asymmetric interactives for the impact of good and bad "news" on the volatility of the variance that proved insignificant.

Table 5 presents our most parsimonious model without and with a time-break dummy variable in the above regression: AR(1) GARCH process for the variance, and only the first lagged squared residual for the ARCH variable. Both ARCH and GARCH effects are significant at the 1% level in both columns. Plot 2 displays the graph of the variance with a period dummy variable only since the alternative shows little change. The evidence of the volatility of the on-offshore differential rate adds an instability dimension of the mean-reverting stationarity of the dual rate; both dimensions of stability are relevant to policy deliberation of full exchange rate unification. We also note the evidence of stationarity for the on-offshore rates reported in some studies are based on first-differenced exchange rate series; the evidence of the differential rate tells a different story.

**Plot 2 Variance from the residual of the regression of $\Delta ln\_HR_t$ on $\Delta ln\_YR_t$ and $D_t$,**



## VI. Conclusions

The focus of this paper has been to obtain some preliminary evidence for the predictive power of an approach based on the differential exchange rate to modeling dynamics behavior of China's parallel market exchange rate. The evidence suggests offshore and onshore rates are cointegrated, thereby opening the door to an assessment of the predictive power of the exchange rate differential as an error correction term in VEC models. We presented linear and Markov chain two-equation VEC models, although containing some of the short-term dynamic lag effects, most are due to the EC term. Our results favor the two-state Markov chain estimates that allow for regime change and offer transitional probabilities; highly persistent for the regression $\Delta ln\_YR_t$ series in regime two. We also examined briefly the volatility of the residual variance, which we found. to be nonstationary, and presented evidence that the variance is significantly affected by first lag of ARCH and GARCH processes. The stationarity of the exchange rate differential series appear to favor a move toward full unification; however, the volatility of the differential rate squared residual series suggests more stabilization policies are necessary ahead of such a major decision.

A key area not yet addressed is the forecasting performance of the differential rate approach compared with those more prevalent in the literature based on fundaments such as growth, inflation, money supply, and other macroeconomic factors. Our evidence must be regarded as preliminary; more critical evidence should come from a comparison of the forecasting accuracy of the differential rate and fundamental. based approaches to short and long term exchange rate dynamics in –exchange rate parallel markets.

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| Table 1 Dickey-Fuller Tests of Stationarity of $ln\_BR_t$ and $ln\_OR_t$* | | | | | | |
|---|---|---|---|---|---|---|
| | No. of lags | Test-statistic | 1%critic. val | 5%critic. val | 10%critic.val | app. *p*-value |
| $ln\_HR_t$ | 1 | -2.020552 | -3.451214 | -2.87062 | -2.571670 | 0.2780 |
| $ln\_YR_t$ | 1 | -1.978617 | -3.451214 | -2.87062 | -2.571670 | 0.2963 |
| Cointegration | 0 | -5.113 | -3.451 | -2.871 | -2.572 | 0.0000 |

*MacKinnon one-sided *p*-values

| Table 2 Test of nonlinearity based on the null linear VEC (absolute *t*-ratios in brackets) | | | |
|---|---|---|---|
| Dep. Var: $\Delta ln\_HR_t$ | | | |
| | | Linear *VEC* | Polynomial*VEC* |
| | | Coefficient. Estimates | Coefficient. Estimates |
| C | | -0.000003.19) | -0.000010(0.47) |
| E(-1) | | -0.099706(1.96) | -0.096367(1.37) |
| ΔLG_HR(-1) | | 0.133385 (1.40) | -0.046317(0.29) |
| ΔLG_YR(-1) | | 0.255164(2.05) | 0.402419(2.19) |
| M1: (ΔLG _HR(-1)$^2$) | | | 119.2657(3.00) |
| M2: (ΔLG _HR(-1)*E(-1)) | | | 0.338704(1.71) |
| M3:(ΔLG _HR(-1)ΔLG _YR(-1)) | | | -155.3793(3.40) |
| M4: (ΔLG _H(-1)$^3$) | | | 12426.79(2.70) |
| M5: (ΔLG_HR(-1)$^2$*E(-1)) | | | 120.4867(0.06) |
| M6: (ΔLG _HR(1)$^2$*ΔLG _YR(-1)) | | | -18939.72(3.05) |
| M7:(ΔLG_HR(-1)$^4$) | | | -17324.10(3.90) |
| M8:(ΔL_HR(-1)$^3$*E(-1)) | | | -535294.2(2.40) |
| M9:ΔL_HR(-1)$^3$*ΔL_YR(-1)) | | | 2646310. (4.05) |
| Null hypothesis *Ho:M1=M2=M3=M4=M5=M6=M7=M8=M9=0* | | | |
| Test Statistic | Value | Df | Probability |
| F-statistic | 2.06905 | (9, 370) | 0.0314 |
| Chi-square | 18.6214 | 9 | 0.0286 |
| Dep. Var: $\Delta ln\_YR_t$ | | | |
| | | Linear *VEC* | Polynomial*VEC* |
| | | Coefficient. Estimates | Coefficient. Estimates |
| C | | -0.00005 (0.39) | 0.000210(1.33) |
| E(-1) | | 0.073717(1.05) | -0.048338(1.00) |
| ΔLG_HR(-1) | | 0.244382(2.68) | 0.018312(0.20) |
| ΔLG_YR(-1) | | | 0.278897(2.20) |
| O1: (ΔLG_YR(-1)$^2$) | | | 15.54551(0.37) |
| O2:(ΔLG_YR(-1)*E(-1)) | | | -34.48435(1.72) |
| O3:( ΔLG_YR(-1)* ΔLG_HR(-1)) | | | -0.334748(0.01) |
| O4: (ΔL_YR(-1)$^3$) | | | -5264.106(0.86) |
| O5: (ΔL_YR(-1)$^2$*E(-1)) | | | -2663.403(0.71) |
| O6:( ΔL_YR(-1)$^2$*ΔL_HR(-1)) | | | 3156.146(0.51) |
| O7: (ΔL_YR(-1)$^4$) | | | -565622.1(0.64) |
| O8: (ΔL_YR(-1)$^3$*E(-1)) | | | 981694.2 (1.84) |
| O9:( ΔL_YR(-1)$^3$*ΔL_HR(-1)) | | | 769222.9(1.10) |
| Null hypothesis *Ho:O1=O2=O3=O4=O5=O6=O7=O8=O9=0* | | | |
| Test Statistic | Value | Df | Probability |
| F-statistic | 0.9207 | (9, 370) | 0.5070 |
| Chi-square | 8.28650 | 9 | 0.5056 |

| Table 3 Two-state Markov Switching Estimation of weekly off-onshore exchange rate 2010. to 2016., (abs. *z*-values in brackets) | | | | | | |
|---|---|---|---|---|---|---|
| *Dep. Var. : $\Delta ln\_HR_t$* | | | | | | |
| | Regime 1 | Regime 2 | Transit. $P_{11}$ | Transit. $P_{22}$ | Duration $P_{11}$ | Duration $P_{22}$ |
| *C* | -0.000508 (1.818949) | 0.000686 (1.548) | *0.4383* | 0.3129 | 1.78 | 1.46 |
| $E_{t-1}$ | -0.0479289 (5.032) | 0.296715 (4.190) | | | | |
| $\Delta ln\_HR_{t-1}$ | -0.225459 (0.949) | 0.295539 (2.292) | | | | |
| $\Delta ln\_YR_{t-1}$ | 0.205298 (0.788) | 0.054262 (0.248) | | | | |
| *Dep. Var. : $\Delta ln\_YR_t$* | | | | | | |
| | Regime 1 | Regime 2 | Transit. $P_{11}$ | Transit. $P_{22}$ | Duration $P_{11}$ | Duration $P_{22}$ |
| C | 0.009232 (9.783) | -0.000260 (2.23) | 0.1311 | 0.9744 | 2.5 | 99.45 |
| E(-1) | -0.378001 (3.381) | -0.064183 (1.51) | | | | |
| Δln_HR(-1) | -1.943177 (6.662) | 0.078567 (1.24) | | | | |
| Δln_YR(-1) | 2.648705 (7.221) | 0.160140 (2.007) | | | | |

| Table 4- LM tests of ARCH Effects in the squared residuals ($Et^2$) of the Exchange Rate Differential eq. (2) with and without a time dummy* | | |
|---|---|---|
| Dep. Var.: $E_t^2$ | Without Break Dummy | With Break Dummy |
| | Coefficient Estimates | Coefficient Estimates |
| $(E_{t-1})^2$ | 0.720559 (18.440) | 0.702181 (17.444) |
| *C* | 0.000004 (1.941) | 0.000004 (1.982) |
| *LM Ho*: $\beta_Y$=0 | Chi2: 162.935, df:1, Prob.Chi2>0.0000<br>F-Stat:340.031, Prob. F(1, 309): 0.0000 | Chi2: 154.304, df:1, Prob. Chi2>0.0000<br>F-stat:304.282, prob. F(1, 309):0.0000 |

*estimates obtained from the residual of a regression of $\Delta ln\_BR_t$ and $\Delta ln\_OR_t$; without a drift, and with and without a time-break dummy.

| Table 5 ARCH and GARCH Estimates of Volatility of the Residual Variance* (ab. z-values in brackets) | | |
|---|---|---|
| | Without Time-Break Dummy | With Time-Break Dummy |
| $ARCH_{t-1}$ | 0.856475 (7.254) | -0.0938919(8.223) |
| $GARCH_{t-1}$ | 0.301194 (6.863) | -0.149462(4.360) |
| *C* | 0.0000004 (4.379) | 0.0000006 (6.745) |

*estimates obtained from the residual of a regression of $\Delta ln\_BR_t$ and $\Delta ln\_OR_t$ without a drift; with and without a time-break dummy.