The Short and Long run Dynamics of Iran's Dual Exchange Rate

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*Abstract*: The unification of Iran's dual exchange rates and the stability of its foreign exchange rates have been key policy concerns to many, both, inside, and outside, the government. This paper explores the short-term, and long-term, dynamics of the relationship between the time-series of Iranian black-market and official exchange rates using 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 for a linear two-equation vector error correction (VEC) model against nonlinearity, and further estimate a three-regime Markov-switching chain model to evaluate the predictive power of the exchange rate differential for the behavior of Iran's exchange rates. We find that the one-period lagged differential displays statistically significant short-term disequilibria correction effects toward the long-run equilibrium exchange rate paths. The Lagrange Multiplier (LM) tests for ARCH effects in the regression residual of the black market exchange rate on the official rate suggest volatility, and report significant estimated ARCH, GRACH effects. We then examined the volatility of the two series and find significant threshold GARCH effects on the variance of the residual. The study suggests the exchange rate differential approach can provide an important predictive/forecasting tool in Iran's exchange rate policy formation and demonstrate the feasibility of a full unification of its dual exchange rates.

## I. Introduction

The literature on exchange rate analysis in parallel markets has relied extensively on the autoregressive effects of the black market and official 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) find 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; hence, 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. 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 addresses this gap. It demonstrates how the incorporation of the differential rate approach into the 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 linear and Markov-switching vector error correction (VEC) models based on the differential rate. We also further examine the volatility of the residual variance obtained from the exchange rate differential by ARCH dynamic processes. Such an analysis has policy relevance to many economies that function with parallel exchange rate markets. In this paper, we examine this approach with an application to Iran, a country with a long history of a dual exchange rate, to demonstrate a framework for modeling the interaction of the short-term and long-term exchange rate dynamics. We apply that approach to a monthly time-series of the black market and the official exchange rate for Iran, covering 1980: 1 to 2012: 1. Our results support the existence of a long-run co-integrated relationship between the two exchange rates, and our estimates of linear and three-regime Markov VEC models indicate a significant, one unit-time, predictive role of the differential variable on the exchange rate time-series. Moreover, our LM tests reveal the non-stationary nature of the second moment of the differential residual, and the presence of significant ARCH effects; and the estimates of ARCH and GRACH processes indicate reduced volatility. This is a preliminarily and partial study of Iran's exchange rate time-series based on the differential rate approach, intended to be developed further for comparison of the forecasting error of the differential-rate-based approach with the fundamental-based approaches.

Section 1 reviews exchange rate time-series literature that bear directly or indirectly on the parallel currency markets; section 2 sets out the statistical tests and econometric model specifications relevant to a differential rate approach, and section 3 is a brief discussion of the data sets. Section 4 presents the outcome for implementing the econometric approach outlined in section 2, and a final conclusion sums up the key pints of this study and some caveats.

## II. Review of Literature

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 the expected speculative profit, and the official exchange rate, 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 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 conclusion. Much of these later developments follow the approach of analyzing the black market exchange rate as a function of fundamentals such 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 cointegrated relationships between the exchange rate and fundamentals and superior forecasting accuracy compared with a random walk based on their panel time-series for a sample of developed economies; 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, with a much lower degree of cross-sectional correlation, and reported similar results. Malone and Ter Horst (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 less extent, the official exchange rates, are the two most important determinants of the black market 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 Cerra and Saxena, test for co-integration of the exchange rate with the fundamentals, and employed forecasting models with deviations of the exchange rate from the fundamentals, and their lagged values, as independent variables.

This study examines a relatively novel alternative approach to the determinant of the exchange rate in a parallel currency based on the information contained in the exchange rate series themselves. As Information about the fundamentals with significant influence on black market exchange rate 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 exchanger rate gap tends toward zero provides a test of the cointegration of the two rates, see Engle and Granger (1987). Provided co-integration is confirmed, a one-period deviation of the black market rate from the official rate, that is, the lagged error term of the regression of the black market rate on the official rate, corrects for short term divergence from equilibrium. The error correction model presents an econometric model for long-run dynamics with a short-term disequilibrium error term embedded within the model as the key determinant of correction toward the 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 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) 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

[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 by price rather quantity change, notably in food markets, see Koohi-Kamali (2012).

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 compared 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. Volatile time series have leptokurtic distributions with fatter tails than the normal, and 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 normality. 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 periods of sharp policy change. 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. It finds relaxation of the official rate increases 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 are monthly. 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. In this case, 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. Maziad and Kang’s (2012) application of the ARCH processes to the Chinses onshore and offshore exchange rates of volatility do allow for the spill-over effects between the two rates, but that study too is not an application of ARCH processes to the time-series of the exchange rate differential. Li (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. The neglect of this issue is all the more important as the evidence on the ability of the ARCH effect terms to forecast volatility with reasonable accuracy should be of major interest.

We selected Iran for an application of the above approach to a dual exchange rate economy because the empirical evidence obtained would have important policy implications, especially regarding her exchange rate unification plan. The literature on Iran has highlighted the positive correlation between the black market rate and inflation. The Pesaran (1997) study for 1980-88 demonstrated that, when Iran's domestic prices have increased more than those of its trading partners, the black market exchange rate followed a similar upward move, and has provided evidence that Iran's persistent inflation is a key determinant of the rising price of international currencies (in Rials). Pesaran's estimates suggest no less than 40% of Iran's exports were priced at black market rates. The Valadkhani (2004) study covers the 1960-2002 period, and confirms the link between high inflation and high black market exchange rate in Iran. It suggests that for every percentage increase in the log of the price ratio, the black market exchange rate increased by 2.5% over that prevailing in 1960-2002. Based on these results, this study reported estimates for a single-equation error correlation model with the log of current black market exchange rate regressed on the first differences in the log of the CPI ratio and the log of GDP, both lagged one period, plus an error correction term estimated to equal to -.20 (t-ratio=1.8), i.e. therefore a 20 percentage point of divergence from its long-run mean value is eliminated in every one year period), see below. The application of a linear error correction model is hard to justify for the entire 1960-2002 period. Tavakolian and Ibrahimi (2012) test this hypothesis regarding the behavior of the Central Bank of Iran (CBI) policy over 19888 to 2009 t with respect to the stabilization of Iran's exchange rate, and reduction of its volatility; by estimating a two-state Markov switching chain model to take into account the official change in the exchange rate policy after 20002. Although the focus of this study is different from the one earlier, the reported evidence also suggests inflation remains an important influence on Iran's conduct of her managed floating exchange rate. The major question not addressed, however, is whether the timing for exchange rate unification is empirically supported. A cointegration test based on the exchange rate differential and a test of volatility of that rate provide a part of the answer to that question.

## III. Econometric Model Specification

In this section we outline VEC model specifications that are tied to the basic characteristics of Iran's dual exchange rate. Consider the following ARIMA model for the basic relationship between the black market exchange rate (BR) and the official exchange rate (OR) from which the ECM can be derived by imposing the parameter restrictions implied by that model.

$$ln\_BR_t = \beta_0 + \beta_1 ln\_OR_t + \beta_2 ln\_OR_{t-1} + \beta_3 ln\_BR_{t-1} + e_t \qquad (1)$$

Solving for long-term equilibrium ($ln\_BR_t = ln\_BR_{t-1}$; $ln\_OR_t = ln\_OR_{t-1}$) lead to the ECM

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

$e^*_{t-1}$ in (2) represents the one-period error correction (EC) disequilibrium term. The inclusion of a drift term is decided empirically with reference to the graph of the differential rate given in plot 1 that displays fluctuations around non-zero mean(s). This suggests the inclusion of a drift term $\phi$ for the specification of the $e^*_{t-1}$ equation[2].

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

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

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

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

[2] Phyliktis and Kassimatis (1994) advocate a differential rate exclusive of a drift term arguing that Dornbusch *et. al*. (1983) stock-flow mode implies an equilibrium black market premium rate defined by the black market *spread* excluding a drift term. This specification is obviously not supported in plot 1.

$\theta_i$ for $i$= 1, 2 is a measure of 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 official and black market rates can be adequately modeled by the linear VEC model in (4) and (5). Such an approach faces a parameter identification problem, for example, when each regime is defined by a 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 non-linearity are based on the null hypothesis of a linear model against the alternative nonlinear model, which nests the linear specification. The literature, especially on smooth transition models, recommends a Taylor expansion around the parameter of the transitional variable, typically the lag of the dependent variable, to test for the additional polynomial and interactives 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 VECM of (4) and (5) specified as

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

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

where $s_{1t}$ = $ln\_BR_{t-1}$ and $s_{2t}$ = $ln\_OR_{t-1}$ 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. The first-order Markov regime-switching assumes that the transitional probabilities of regime change depend only on the move from the previous regime, namely,. moving from a recession period to that of expansion, is independent of how long the recession has lasted. They are constant for moves from $i$ to $j$ regimes and 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-reversion 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 for volatility. The following is the simplest version of such an equation selected by Augmented Dicky-Fuller (ADF) tests as the model that has white-noise error.

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

where $e_t$ is the disturbance term obtained from the regression of $ln\_BR_t$ on $ln\_OR_t$ (with a drift), and $h_t$ stands for its variance; with necessary conditions for covariance stationarity $\delta > 0, \alpha_i \geq 0, \beta_i \geq 0, (\Sigma\alpha_i + \Sigma\beta_i) < 0$.

We further test the structural stability of this model by adding two dummy variables that divide the series into the three exchange rate regimes depicted in plot 1, namely $D_1$=1 if date <= 1992 and zero otherwise (observations = 157); $D_2$=1 if 1992: 1<date<=2002:s zero otherwise (observations= 265). Employing the residual series from this extended version of (3), eq. (8) is then re-estimated for ARCH effects.

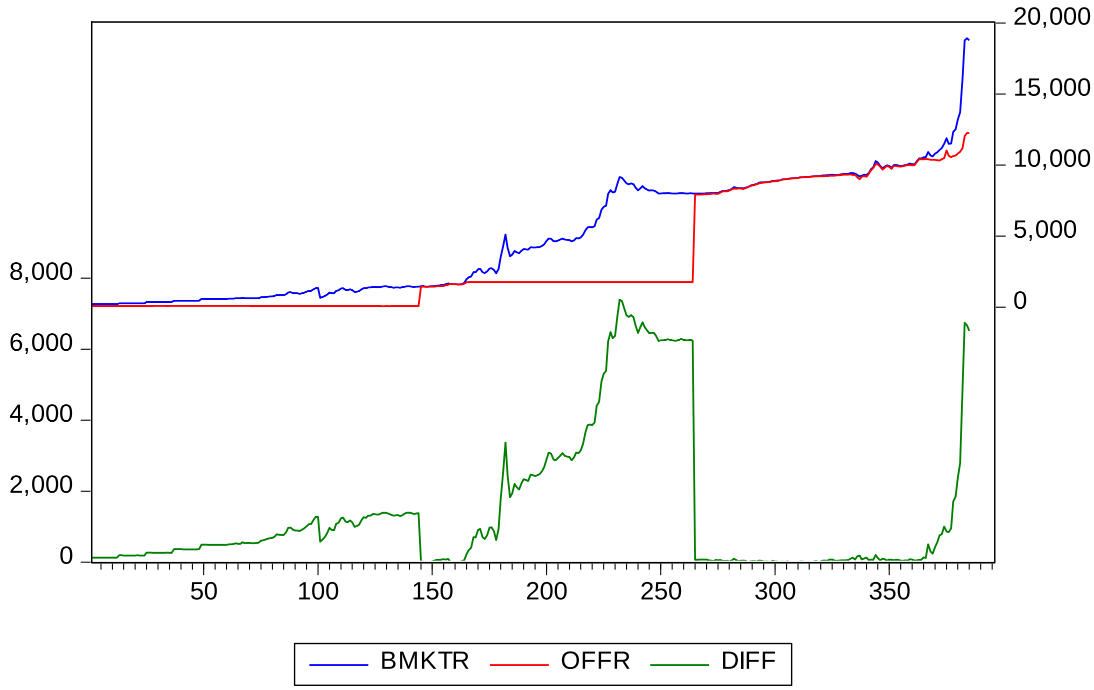
## IV **Data**

Iran unified her multiple official exchange rates into a single one in May 1992; there are two clear exchange rate regimes before and after that date. In fact over the period since 1983, there have been three distinct exchange rates regimes as is evident from plot 1: prior to 1993 there were several fixed exchange rates applied to various categories of goods and services; all simplified into a single fixed rate from March 1993, and finally in March 2002 the Central Bank of Iran (CBI) adopted a policy unified, floating but managed exchange rate, namely, during periods of sharp oil revenue increases in dollar terms, the CBI intervenes in the foreign currency market to ensure the *Rial* appreciates, and in periods of decreases in revenue, it intervenes to ensure depreciation, so as to maintain a constant ratio of dollar holdings to the monetary base. Full unification is currently under consideration.

The data for Iran's exchange rate covers 32 years of monthly series from March 1980-March 2012) , providing 384 observations, and is available online from the CBI. The data set of both Iran's official exchange rate and black-market rate are from the time period between March 1980 and March 2012, during which the Iran's economy underwent both economic boom, and recession. The monthly dimension of the data sets is not without interest, for volatility tests since the temporal aggregation hypothesis suggests a reduction of volatility with a fall in time frequency of the observations.

Plot 1displays the monthly time-series of Iran's dual exchange rate over 1980to 2012A clear pattern of three different exchange rate regimes is evident; moreover, prior to 2002:, the black market rates tend to rise relative to the official rate in both periods, up to 1992,and from 1992to 2002This trend is particularly pronounced over the latter period. However, we note that while the first and last periods display co-movement, the middle period appears to be uncorrelated, a pattern that will affect some of the results presented below.

***Plot* 1** Monthly time-series of Iran's dual exchange rate
(in levels) 1980:1 to 2012:1



## V. Results

We employ the augmented Dickey-Fuller (ADF) test; the Schwartz criterion selects three lags for the black market rate, and one for the official rate, to remove all autocorrelation effects from the disturbance term. The results, reported table 1, indicate the black-market test statistic (-1.737) is not significant at the 10% DF critical value; the official rate (-0.821) is also not significant at 10% critical value. Therefore, both series remain nonstationary, *I* (0); also evident from insignificant *p*-values. Rather than moving forward with *I* (1) exchange rate variables in first differences, we test for the stationarity of their linear combination.

To test for co-integration, we regress the black-market rate on the official rate series, a time-trend, and the first lags of both series suggested by the Schwartz model selection criterion. The ADF test statistic (-2.65), in the last row of table 1, is significant at the 10% DF critical value; a significant *p*-value of 0.083 now indicates a white-noise disturbance term. Since, by the Granger representation theorem, Engle and Granger (1987), the relationship between the two rates has an equivalent error correction representation; we apply a VEC model to estimate the long and short-term dynamics of that relationship.

We first estimate a linear two-equation VEC model, mainly, testing the null hypothesis against the hypothesis of nonlinear VEC processes. The model selection rule picks a VEC model with only the first-order lagged 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, as reported in the first column of table 2. The error correction coefficient estimates are statistically significant at the 5% level, 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 $\Delta ln\_BR_t$ by 1.3 percent and raises $\Delta ln\_OR_t$ by 3.2 percent to bring the two rates closer to the long-term equilibrium path. We also note that only $\Delta ln\_BR_{t\text{-}1}$ in (3) has a significant lagged impact. However, as discussed above, a linear specification of the VEC model has its limitations.

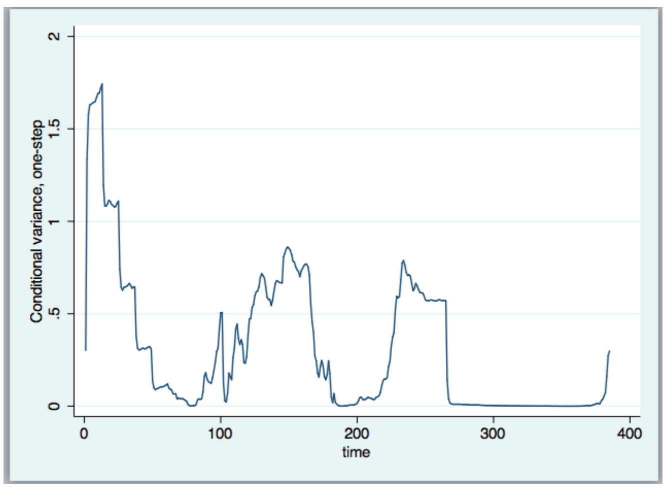
We test the linear VEC specification with a single exchange rate regime against nonlinearty that arises from having changing regimes. A transitional function from one state to another is defined as a function of the transitional variable; since only the first lags are significant, we take $\Delta ln\_BR_{t\text{-}1}$ and $\Delta ln\_OR_{t\text{-}1}$ as the transitional variables, and expand the linear VEC around zero of the slope parameters. We expand (4) and (5) by the cubic terms for the transitional variables together with 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 (18.62) for $df$=9 and $F_{(9, 370)}$ test statistic (2.10) both reject the linear VEC specification for eq. (4) at 5 %, but not for eq. (5), see the second column of table 2, last row. Therefore, at least for eq. (4), a linear VEC is a mis-specification.

Table 3 presents our nonlinear estimates corresponding to the above two-equation VEC system by a three-state Markov-switching chain process specified in (6) and (7).We note that in eq. (6), there are statistically significant error correction terms ($E_{t}$-1) in regimes one and three, but not in regime two. Transitional $p$-values for regime two indicate high persistence; that is, there is high probability that exchange rates in that regime will be followed by another from the same regime. This is reflected in the constant expected duration of transition for eq. (6) reported in the last three cells in the first row. These are obtained from 1/(1-$P_{ii}$) by using the transitional $p$-values from the table for $i$=1, 2. 3: 1.12, 14.64: and 1.55 months. As for eq. (7), there are EC effects significant at 1% in all three states, with constant expected duration as : 1, 47.24, and d 1 month, once again the duration from regime two transition shows high persistence. In all cases, however, there are also significant lagged effects for both rates in all three states. In short, the EC term drives the series toward equilibrium in combination with $AR$(1) effects; high probability observation from regime two to be followed by another from the same regime.

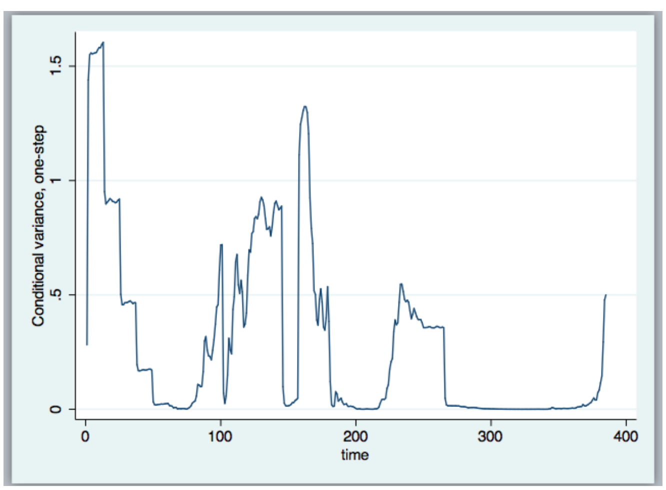
Next, given that the disturbance from (3) has a stationary mean, we examine the volatility of its second moment. Table 4 reports the Lagrange Multiplier (LM) test for ARCH effects in by the first-order autoregessive for the square of the disturbance term obtained from the regression of $\Delta ln\_BR_t$ on $\Delta ln\_OR_t$ , with an intercept, first for the lag effects of the residual obtained from eq. (3), and then by the residual obtained from an addition to the above regression of two dummy variables corresponding to the two break points that delineate the three exchange rate regimes (1992:1) and (2002:1). The $\chi^2$ LM test statistics, in the last row, decisively reject the null hypothesis of no $ARCH$ effects in both models for the residual of eq. (3) obtained with and without dummy variables. We estimated several families of $ARCH$ models for the residual variance of (3) 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, two time-break dummy variables: 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; the lagged GARCH obtained with time dummy variables. However, it is only mildly significant. Plots 2 and 3 display the graph of the variance with, and without period dummy variables.

We note the following from our ARCH analysis. First, perhaps it is notable that the observations aggregated to monthly levels still reveal volatility, given the prediction of the temporal aggregation hypothesis. Second, even though the differential gap is mean-stationary, its variance is not, which has implication for exchange rate unification discussed below.

Plot 2: variance of eq. (8) without dummy variables




Plot 3: variance of eq. (8) with dummy variables



## V. Conclusion

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 Iran's parallel market exchange rate. The evidence suggests the black market and official rates are co-integrated; therefore, hence opening the door to an assessment of the predictive power of the exchange rate differential as an error correction term in a two-equation VEC model. We presented linear and Markov chain VEC models, all of which display statistically significant EC effects, though there are also significant effects due to lagged variables in the Markov estimates. Our results for the three-state Markov estimates allow for exchange rate regime change and offers transitional probabilities for regime change effects that suggest high persistence for regime two. We also examined briefly the volatility of the residual variance, found it to be nonstationary, and presented evidence that the variance is significantly affected by the first lag of ARCH and GARCH processes. We note that our ARCH analysis is based on a series cointegrated in levels, not in first-differences. Our results on the stability of the differential rate, provides partial answers as to whether Iran is ready for full exchange rate unification. However, successful unification also requires the relative absence of volatility.

Finally, a key area not yet addressed is the forecasting performance of the differential rate approach compared with the ones more prevalent in the literature based on fundamentals 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 forecasting accuracy of the differential rate and fundamentals-based approaches to short and long term exchange rate dynamics in parallel markets.

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| Table 1 Augmented 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 |
| $ln\_BR_t$ | 3 | -1.911 | -3.450 | -2.875 | -2.570 | 0.321 |
| $ln\_OR_t$ | 1 | -0.821 | -3.449 | -2.875 | -2.570 | 0.1466 |
| cointegration | 1 | -2.745 | -3.449 | -2.875 | -2.570 | 0.0833 |

*MacKinnon one-sided *p*-values

| Table 2 Test of nonlinearity based on the null linear VEC (absolute *t*-ratios in brackets) | | | |
|---|---|---|---|
| Dep. Var: $\Delta ln\_BMR_t$ | | | |
| | | Linear *VEC* | Polynomial*VEC* |
| | | Coef. Estimates | Coef. Estimates |
| C | | 0.017670(3.37) | 0.008160(2.41) |
| E(-1) | | -0.012838(2.14) | -0.013364(2.10) |
| ΔLG_BMR(-1) | | 0.106481 (2.09) | 0.360365(2.84) |
| ΔLG_OFR(-1) | | -0.007548(0.42) | -0.007852(0.32) |
| M1: (ΔLG _BMR(-1)$^{2}$) | | | 0.310377(0.30) |
| M2: (ΔLG _BMR(-1)*E(-1)) | | | -0.008142(0.03) |
| M3:(ΔLG _BMR(-1)* ΔLG _OFR(-1)) | | | 0.158151(0.02) |
| M4: (ΔLG _BMR(-1)$^{3}$) | | | -0.693525(0.19) |
| M5: (ΔLG_BMR(-1)$^{2}$*E(-1)) | | | -2.456195(0.81) |
| M6: (ΔLG _BMR(1)$^{2}$*ΔLG _OFR(-1)) | | | 39.37330(0.24) |
| M7:(ΔLG_BMR(-1)$^{4}$) | | | -6.683606(0.42) |
| M8:(ΔL_BMR(-1)$^{3}$*E(-1)) | | | 8.837051(0.60) |
| M9:ΔL_BMR(-1)$^{3}$*ΔL_OFR(-1)) | | | -431.2718(0.45) |
| 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: $\Delta ln\_OFR_t$ | | | |
| | | Linear *VEC* | Polynomial*VEC* |
| | | Coef. Estimates | Coef. Estimates |
| C | | 0.013370 (1.46) | 0.012167(1.18) |
| E(-1) | | 0.031820 (1.86) | 0.032975(1.58) |
| ΔLG_BMR(-1) | | -0.006421(0.04) | -0.070295(0.25) |
| ΔLG_OFR(-1) | | | 1.458313(1.52) |
| O1: (ΔLG_OFR(-1)$^{2}$) | | | 16.31563(0.45) |
| O2:(ΔLG_OFR(-1)*E(-1)) | | | 2.900547(1.75) |
| O3:( ΔLG_OFR(-1)* ΔLG_BMR(-1)) | | | 5.580162(0.30) |
| O4: (ΔL_OFR(-1)$^{3}$) | | | -36.33348(0.04) |
| O5: (ΔL_OFR(-1)$^{2}$*E(-1)) | | | -22.10401(0.27) |
| O6:( ΔL_OFR(-1)$^{2}$*ΔL_BMR(-1)) | | | 147.8763(0.20) |
| O7: (ΔL_OFFR(-1)$^{4}$) | | | 10.48706(0.01) |
| O8: (ΔL_OFFR(-1)$^{3}$*E(-1)) | | | -13.25121(0.01) |
| O9:( ΔL_OFR(-1)$^{3}$*ΔL_BMR(-1)) | | | -6999.891(0.59) |
| 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 3-State Markov switching chain VEC estimates of Black market & Official monthly exchange rates (absolute *z*-values in brackets) | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Dep. Var: $\Delta ln\_BR_t$ | | | | | | | | | |
| | Regime 1 | Regime 2 | Regime 3 | trans. $P_{11}$ | trans.$P_{22}$ | trans. $P_{33}$ | dur$P_{11}$ | dur$P_{22}$ | dur$P_{33}$ |
| *C* | -0.085582 (10.148) | 0.004804 (3.617) | 0.117635 (14.104) | 0.1070 | 0.9317 | 0.3329 | 1.120 | 14.64 | 1.499 |
| $E_{t-1}$ | -0.409673 (33.861) | 0.000384 (0.153) | -0.043647 (2.259) | | | | | | |
| $\Delta ln\_BR_{t-1}$ | 0.050340 (0.613) | -0.022359 (1.002) | 0.306977 (4.373) | | | | | | |
| $\Delta ln\_OR_{t-1}$ | -8.699 (18.941) | 0.001550 (0.226) | 0.279930 (0.818) | | | | | | |
| Dep. Var: $\Delta ln\_OR_t$ | | | | | | | | | |
| | Regime 1 | Regime 2 | Regime 3 | tansi. $P_{11}$ | trans.$P_{22}$ | trans. $P_{33}$ | dur$P_{11}$ | dur$P_{22}$ | dur$P_{33}$ |
| *C* | 0.032457 (2.725) | 0.001266 (2.027) | 0.426828 (22.289) | 0.00000 | 0.9788 | 0.000031 | 1.000 | 47.24 | 1.000 |
| $E_{t-1}$ | 0.377058 (4.612) | -0.003716 (3.222) | 1.588562 (64.951) | | | | | | |
| $\Delta ln\_BR_{t-1}$ | -0.897745 (2.919) | 0.009917 (1.015) | 18.62629 (14.248) | | | | | | |
| $\Delta ln\_OR_{t-1}$ | 2.033297 (3.210) | 0.001300 (0.379) | 57.53138 (62.675) | | | | | | |

| Table 4- LM tests of ARCH Effects for the Exchange Rate Differential | | |
|---|---|---|
| | Without Break Dummies | With Break Dummies* |
| Dep. Var.: $E^2_t$ | Coefficient Estimates | Coefficient Estimates |
| $E^2_{t-1}$ | 0.542091 (211.39) | 0.243540 (36.08) |
| C | 4.217572 (178.7) | 5.396352 (172.16) |
| *LM Ho*: $\beta$ of $E^2_{t-1}$=0 | Chi2-stat: 368.62, Prob. Chi2 (1): 0.0000<br>F-stat: 9768.67, Prob. F(1, 381): 0.0000 | Chi2-stat: 354.79, Prob. Chi2(1): 0.0000<br>F-stat: 4791.01, Prob. F(1, 381): 0.00000 |

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

| Table 5 ARCH and GARCH Estimates of Volatility of the Eq. (2) Residual Variance (ab. z-values in brackets) | | |
|---|---|---|
| | Without Time-Break Dummies | With Time-Break Dummies* |
| $ARCH_{t-1}$ | 0.887872(3.787) | 1.095655 (5.511) |
| $GARCH_{t-1}$ | 0.180061 (3.838) | 0.080371 (1.536) |
| *C* | 0.0000005 (1.014) | 0.0000009 (1.011) |

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