The Rial-Dollar Exchange Rate and Purchasing Power Parity Theory

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Abstract

Since the 1979 Islamic Revolution in Iran, Iranian currency has depreciated from 70 rials per dollar to as low as some 36000 rials per dollar. Has this movement followed the path predicted by the well-known Purchasing Power Parity (PPP) theory? In this paper we show that the answer is in the affirmative and the dominating factor causing the decline is domestic inflation. Following the theory, we predict a rate of almost 47000 rials to the dollar.

Introduction

A review of the value of the Iranian rial against the U.S. dollar reveals that it has continuously lost value and mostly since the the 1979 revolution. The free market exchange rate reveals that from 1933 until 1979, the dollar rose from 11.20 rials in 1933 to 98 rials in January 1979. However, the rate as of early 2013 stood at 36,300 rials per dollar.\(^3\) This paper uses the economic theory of purchasing power parity (PPP) to explain this abnormal rise in the rial-dollar exchange rate. If we consider the dollar to be like any other commodity, the price of land, housing, food, television sets and, cars, has risen, why not the price of the dollar? Accordingly then the main explanation for the rise in the dollar must be domestic inflation. Thus, whatever has contributed to domestic inflation in Iran, has also contributed to the increase in the value of the dollar.\(^4\) The PPP theory theory explains the link between the exchange rate and relative prices.

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\(^3\) The rate is at the closing on January 31, 2016.

\(^4\) Such factors include the Iran-Iraq war, loss of oil revenue, sanctions by the United Nations and the West, and
This theory is outlined in section II and applied and tested using monthly price data in the post-revolutionary period from January 1979 to July 2015. The results are summarized in Section III.\textsuperscript{5} Test applied to the level of both variables as well as to their first-differences are reported in Table 1. It is clear that regardless of the number of lags used in the ADF test, only first-differenced variables are stationary. Therefore, both \( \ln \text{EX} \) and \( \ln (\text{PIR}/\text{PUS}) \) are integrated of order one or \( I(1) \). Now we must show that a linear combination of the two variables proxied by \( \varepsilon_t \) is integrated of order zero or \( I(0) \). To that end we estimate equation (2) and apply the ADF test to the residuals. The results are as follows:\textsuperscript{8}

\[
\hat{\alpha} = 9.63 \quad (565.2) \quad \hat{\beta} = 0.842 \quad (107.6) \\
\overline{R}^2 = 0.96 \quad \text{ADF}(10) = -2.61
\]

If the PPP is to hold, we would expect the estimate of slope coefficient in (2) to be one which it is not. We would also expect the residuals to be stationary. The ADF test statistics of 2.61 is less than its critical value of 3.76 in absolute value, implying that the residuals are not stationary, thus rejecting the PPP theory.\textsuperscript{9}

By assuming the slope coefficient in equation (2) to be one, we also assume that the inflation rate in Iran and inflation rate in the US have a similar impact on the exchange rate, though in opposite directions. Clearly, this cannot be the case since very little trade took place between the two countries during our period of study. To test this hypothesis, we separate the two terms and estimate the following specification:

\[
\ln \text{EX}_t = a + b \ln \text{PIR}_t + c \ln \text{PUS}_t + \varepsilon_t 
\]

\textbf{The Purchasing Power Parity (PPP) And Testing Method}

Let \( \text{EX} \) denote the exchange rate between rial and the U.S. dollar. The purchasing power parity theory (PPP hereafter) basically claims that in the long run the exchange rate must be equal to the price ratio of the two countries.\textsuperscript{6} Denoting the price level in Iran by \( \text{PIR} \) and the price level in the U.S. by \( \text{PUS} \), the PPP could be outlined by equation (1):

\[
\text{EX} = \frac{\text{PIR}}{\text{PUS}} 
\]

In order to see if the two variables follow each other, we plot each of them in Figure 1.

|Figure 1 here|

From Figure 1 we see that indeed, the two variables move together most of the time. However, the relation could be spurious unless we establish cointegration between the two variables. If we follow the cointegration concept of Engle and Granger (1987), we show that the two variables are integrated of the same order \( d \), but a linear combination of the two is integrated in an order less than \( d \).

A common practice is to express (1) in a log-linear format as in (2):

\[
\ln \text{EX}_t = \alpha + \beta \ln \left(\frac{\text{PIR}}{\text{PUS}}\right)_t + \varepsilon_t 
\]

The results of the ADF test applied to the level of both variables as well as to their first-differences are reported in Table 1. It is clear that regardless of the number of lags used in the ADF test, only first-differenced variables are stationary. Therefore, both \( \ln \text{EX} \) and \( \ln (\text{PIR}/\text{PUS}) \) are integrated of order one or \( I(1) \). Now we must show that a linear combination of the two variables proxied by \( \varepsilon_t \) is integrated of order zero or \( I(0) \). To that end we estimate equation (2) and apply the ADF test to the residuals. The results are as follows:\textsuperscript{8}

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To see why \( \varepsilon_t \) in (2) is a proxy for the linear combination of two variables, all we need to do is to solve (2) for \( \varepsilon_t \).

\textbf{Numbers inside the parentheses for coefficient estimates are absolute values of the t-ratios. However, number inside the parenthesis next to ADF statistic is number of lags in the test. Note that data on price levels come from International Financial Statistics of the IMF. Data on the rial-dollar rate are mostly from Bahmani-Oskooee (2005) and author’s own collection.}

\textbf{Note that we have selected the ADF statistic by minimizing the AIC criterion. However, the ADF statistic was insignificant at all lags (1-12). We also made sure that both \( \ln \text{PIR} \) and \( \ln \text{PUS} \) are \( I(1) \).}
If PPP is to hold, an estimate of $b$ should be one and that of $c$ negative one. Furthermore, $\varepsilon_t$ should be I(0). The OLS estimate of (3) is as follows:

$$\hat{a} = 10.77(7.52) \quad \hat{b} = 0.88(15.6) \quad \hat{c} = -1.13(3.07)$$

$$R^2 = 0.97 \quad ADF(10) = -2.64$$

It is clear from the results that the estimate of $b$ is not one and that of $c$ is not negative-one, although they are close to their expected values. These estimates are significant, since the absolute value of the ADF statistic of the residuals is less than the critical value of 3.76, the null of unit root in the residuals cannot be rejected, hence the residuals are non-stationary, rejecting cointegration.\(^\text{10}\)

The ADF test applied to the residuals of equation (3) is said to suffer from low power. We therefore, shift to an alternative approach of testing for cointegration that incorporates short-run dynamic adjustment of variables. To this end, we follow Coe and Serletis (2002) and Pesaran et al. (2001) bounds testing approach and rewrite equation (3) in an error-correction format as follows:

$$\Delta \ln EX_t = \alpha + \sum_{i=1}^{12} \delta_i \Delta \ln EX_{t-i} + \sum_{i=0}^{12} \delta_i \Delta \ln PIR_{t-i} + \sum_{i=0}^{12} \phi_i \Delta \ln PUS_{t-i} + \lambda_i \ln EX_{t-i} + \lambda_i \ln PIR_{t-i} + \lambda_i \ln PUS_{t-i} + \mu_i \Delta \ln PUS_{t-i} + \mu_i \Delta \ln PIR_{t-i} + \mu_i \Delta \ln EX_{t-i} + \epsilon_t \quad (4)$$

Equation (4) is an error-correction model that is similar to the Engle and Granger (1987) specification. The difference is that rather than including lagged error term from (3) we have included its proxy represented by the linear combination of lagged level variables. In order to justify the inclusion of lagged level variables, Pesaran et al. (2001) propose applying the familiar F test to establish their joint significance as a sign of cointegration. However, they demonstrate that the F test in this application has new critical values that they tabulate. Since these critical values account for integrating properties of variables, under this method there is no need for pre-unit root testing and variables could be combination of I(0) and I(1), though our variables are all I(1).

Equation (4) is estimated after imposing 12 lags on each variable. Following the literature, we use Akaike’s Information Criterion (AIC) to select an optimum model. The selected model was an order of (12, 3, 0) and the normalized log-run coefficients and the F test for joint significance of lagged level variables were estimated to be:

$$\hat{\alpha} = 9.46(0.98) \quad \frac{\hat{\lambda}_1}{\lambda_0} = 0.79(2.10)$$

$$\frac{\hat{\lambda}_2}{\lambda_0} = -0.74(0.30) \quad F = 2.90$$

From these results it is clear that while the Iranian price level carries a significant coefficient, the U.S. price level does not. This implies that the inflation rate in Iran is more relevant and the main determinant of the rial-dollar exchange rate in Iran. However, the F test applied to the joint significance of lagged level variables is insignificant given its upper bound critical value of 4.14 at the 10% level of significance.\(^\text{11}\) However, according to Bahmani-Oskooee and Tanku (2008), there is an alternative way of establishing cointegration. Using the long-run coefficient estimates and long-run model we generate the error term, usually called error-correction term labeled as ECM. We then replace the linear combination of lagged level variables in (4) by ECM\(_{t-1}\) and estimate the resulting specification at the same optimum lags. A significantly negative coefficient for ECM\(_{t-1}\) will support convergence toward long run or cointegration. Once this specification was estimated, the coefficient estimate of ECM\(_{t-1}\) was -0.03 with a t-ratio of -2.87, supporting cointegration.

**Summary And Conclusion**

Since the Islamic Revolution in 1979, the Iranian rial has been under pressure and has lost much of its value. The free market rial-dollar rate has gradually moved from 70 rials per dollar to 36,300 rials per dollar as of February 2013. While most of the pressure in late 2011 and 2012 is attributed to economic sanctions by the U.N., and the U.S. and Chortareas and Kapetanios (2004), and Bahmani-Oskooee and Goswami (2005).

\(^{10}\) For some other studies that have tested the PPP using the black market exchange rate see Bahmani-Oskooee (1993), Phylaktis and Kassimatis (1994), El-Sakka and McNabb (1994), Sanchez-Fung (1999), Kargbo (2003),
other Western countries, we demonstrate that during the last three decades, inflation in Iran has been the main source of the devaluation of the rial.

In this paper we tried to establish the link between the rial-dollar rate and inflation differential between Iran and the U.S., following the purchasing power parity theory. Using monthly data from January 1979 – July 2015, our empirical results reveal that the dominating factor in the decline of the rial is domestic inflation. During the study period, the Iranian Consumer Price Index has moved up from 0.43 to almost 288, an increase of 670-fold. If we were to increase the price of the dollar by this ratio, we would expect a rate of 46,900 rials per dollar. Thus, the current rate of almost 37,000 rials per dollar is almost 20% less than the PPP would predict. Indeed, this is reflected in coefficient estimates of the Iranian inflation rate in model (4) and could be due to imperfections in the PPP or intervention by Iran’s Central Bank.

Figure 1: Plot of the Rial-dollar rate (EX) and Relative Prices (RP)
Table 1: The Result of ADF Test Applied to Level and First-Differenced Variables

<table>
<thead>
<tr>
<th></th>
<th>LnEX</th>
<th>ΔLnEX</th>
<th>LnRP</th>
<th>ΔLnRP</th>
</tr>
</thead>
<tbody>
<tr>
<td>DF</td>
<td>-.99081</td>
<td>-16.74</td>
<td>0.7798</td>
<td>-13.27</td>
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<tr>
<td>ADF(1)</td>
<td>-.98318</td>
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<td>0.3936</td>
<td>-10.13</td>
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<td>ADF(2)</td>
<td>-.97886</td>
<td>-10.11</td>
<td>0.3543</td>
<td>-8.304</td>
</tr>
<tr>
<td>ADF(3)</td>
<td>-.98864</td>
<td>-9.062</td>
<td>0.2987</td>
<td>-8.032</td>
</tr>
<tr>
<td>ADF(4)</td>
<td>-.98646</td>
<td>-8.551</td>
<td>0.3291</td>
<td>-8.256</td>
</tr>
<tr>
<td>ADF(5)</td>
<td>-.97627</td>
<td>-8.085</td>
<td>0.3940</td>
<td>-7.662</td>
</tr>
<tr>
<td>ADF(6)</td>
<td>-.96496</td>
<td>-7.450</td>
<td>0.4013</td>
<td>-6.632</td>
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<tr>
<td>ADF(7)</td>
<td>-.96674</td>
<td>-7.005</td>
<td>0.3608</td>
<td>-6.026</td>
</tr>
<tr>
<td>ADF(8)</td>
<td>-.98151</td>
<td>-7.061</td>
<td>0.3341</td>
<td>-5.745</td>
</tr>
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<td>ADF(9)</td>
<td>-.94171</td>
<td>-6.617</td>
<td>0.3262</td>
<td>-4.942</td>
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<tr>
<td>ADF(10)</td>
<td>-0.95103</td>
<td>-5.354</td>
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<td>ADF(11)</td>
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<td>-5.525</td>
<td>0.1091</td>
<td>-3.147</td>
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<tr>
<td>ADF(12)</td>
<td>-1.0201</td>
<td>-4.834</td>
<td>-0.0388</td>
<td>-3.101</td>
</tr>
</tbody>
</table>

95% critical value for the augmented Dickey-Fuller statistic = -2.8692
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References


