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Nonlinear ARDL Approach and the Demand for Money in Iran

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Abstract

To account for currency substitution, most studies today include exchange rate as a determinant of the demand for money, in addition to income and interest rate. This tradition goes back to Robert Mundell who introduced this notion in 1963. In this paper, we demonstrate that the failure to find a significant relationship between exchange rate and the demand for money could be due to assuming a linear adjustment mechanism among the variables. Once we introduce nonlinearity in the short run as well as in the long run through partial sum concept, we show that currency appreciation or depreciation could affect the demand for money in an asymmetric manner. This is demonstrated using data from Iran.

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1. Introduction

The demand for money is perhaps one of the areas in economics which has received the greatest attention. Identifying variables, other than income and interest rate, testing for integrating properties of the relevant variables, establishing cointegrating properties of the variables, and finally, testing for stability of the money demand function are some of the important aspects of the literature upon which researchers have concentrated.

Concentrating on the first issue of identifying relevant variables, in 1963 the Nobel laureate Mundell (1963) conjectured that in addition to income and interest rates, the exchange rate could certainly be another determinant of the demand for money. Subsequent studies not only tried to provide clear and intuitive explanations for this assumption, but also empirical support for Mundell's conjecture. Very briefly, the depreciation of domestic currency, or the appreciation of foreign currency, raises the domestic currency value of foreign assets held by domestic residents. If this is perceived as an increase in wealth, the demand for money should increase (Arango and Nadiri, 1981). However, when foreign currency appreciates, if there is the expectation of further appreciation, domestic residents may hold more foreign currency and reduce their demand for domestic currency (Bahmani-Oskooee and Pourheydarian, 1990). Therefore, depending upon the strength of the wealth effect versus the expectation effect, the demand for money could move in either direction. Example of studies which have included the exchange rate in the money demand include McNown and Wallace (1992) for the U.S.; Lee and Chung (1995) for Korea, Miyao (1996) for Japan, Chowdhury (1997) for Thailand, and Bahmani-Oskooee and Gelan (2009) for African countries. The findings are mixed at best.

Failure to find support for the significance of the exchange rate in the demand for money could stem from assuming that the short-run dynamic adjustment process follows a linear path. Once we introduce the non-linear adjustment process into the testing procedure, we show that currency depreciation could have a significant effect on the demand for money. To that end, we introduce the model and methods in Section II. In section III we present our empirical results followed by a summary and concluding remarks in Section IV. Data definition and sources are cited in an Appendix.

2. The Model and Methodology

Out of our interest and for demonstrative purpose, we use data from Iran. Therefore, the specification adopted in this paper follows Bahmani-Oskooee (1996) who estimated the demand for money for Iran by including real income as a scale variable, inflation rate as a measure of the opportunity cost of holding money, and the black market exchange rate, as in equation (1):[1]

$$LnM_t = a + bLnY_t + c\pi_t + dLnEX_t + \varepsilon_t \qquad (1)$$

where M is a measure of real M2 monetary aggregate, Y is a measure of real income, π is the rate of inflation, EX is the exchange rate, and ε is an error term. We expect an estimate of b, which measures income elasticity of the demand for money, to be positive. The inflation rate, rather than the interest rate, is included in the specification to measure the opportunity cost of holding money due to the lack of well-developed financial markets in Iran. As Bahmani-Oskooee (1996) argued, historically in Iran real assets, such as land and housing, are the best alternative for holding cash. Therefore, we expect an estimate of c to be negative. As for an estimate of d, as mentioned in the introduction, it could be negative or positive depending on whether the dollar appreciation increases expectations of further appreciation or is perceived as an increase in wealth (i.e., an increase in domestic currency value of foreign assets).[2]

Models such as (1) are labeled long-run models and their estimates by any means will yield only long-run coefficient estimates once cointegration among the variables is established. Since stability of the money demand is another concern, Laidler (1993, p. 175) argued that one source of instability could stem from excluding short-run dynamics from long-run model. To that end, we specifying (1) in an error-correction modeling format as in equation (2):

$$\Delta LnM_t = \alpha + \sum_{i=1}^{n1} \beta_i \Delta LnM_{t-i} + \sum_{i=0}^{n2} \delta_i \Delta LnY_{t-i} + \sum_{i=0}^{n3} \gamma_i \pi_{t-i} + \sum_{i=0}^{n4} \eta_i \Delta LnEX_{t-i}$$

$$+ \rho_0 LnM_{t-1} + \rho_1 LnY_{t-1} + \rho_2 \pi_{t-1} + \rho_3 LnEX_{t-1} + \mu_t \qquad (2)$$

Equation (2) follows Pesaran *et al.*'s (2001) bounds testing approach who offer a one-step procedure in estimating the short-run as well as the long-run effects. The short-run effects are contained in the coefficients attached to the first-differenced variables. The long-run effects are obtained by the estimates of ρ_1-ρ_3, normalized on ρ_0.[3] However, for the long-run estimates to be valid, cointegration must be established. Pesaran *et al.* (2001) propose applying the standard F test for joint significance of the lagged level variables in (2). However, as they demonstrate, this

[1] This specification is followed by many others in the literature.
[2] Note that, although there are neither economic nor political relations between the U.S. and Iran, the U.S. dollar still serves as a reserve currency and dominates almost all international transactions. For the history of the rial-dollar rate see Bahmani-Oskooee (2005).
[3] For the exact normalization procedure see Bahmani-Oskooee and Tanku (2008) in this journal.

F test has a new distribution with new critical values that they tabulate. An upper bound critical value is provided by assuming all variables in a model to be I(1) and a lower bound critical value is provided when all variables are assumed to be I(0). They demonstrate that the upper bound critical values could also be used if some variables are I(1) and some are I(0). Since almost all macroeconomic time-series variables are either I(1) or I(0), there is actually no need for pre unit-root testing under this method.[4]

Previous studies that estimated any of the above models using any method, assumed that exchange rate changes have symmetric effects on the demand for money. To test this hypothesis, we decompose the movement of the Ln EX variable into its negative (the depreciation of the dollar) and positive (the appreciation of the dollar) partial sum as: $LnEX = LnEX_0 + LnEX_t^+ + LnEX_t^-$ where Ln EX^+_t and $LnEX^-_t$ are the partial sum process of positive and negative *changes* in Ln EX. More precisely:

$$LnEX_t^+ = \sum_{j=1}^{t} \Delta LnEX_j^+ = \sum_{j=1}^{t} \max(\Delta LnEX_j, 0), \quad LnEX_t^- = \sum_{j=1}^{t} \Delta LnEX_j^- = \sum_{j=1}^{t} \min(\Delta LnEX_j, 0)$$

We then follow Shin *et al* (2014) and replace Ln EX in equation (2) by Ln EX^+_t and Ln EX^-_t as in (3) below:

$$\Delta LnM_t = a + \sum_{i=1}^{n1} b_i \Delta LnM_{t-i} + \sum_{i=0}^{n2} c_i \Delta LnY_{t-i} + \sum_{i=0}^{n3} d_i \pi_{t-i} + \sum_{i=0}^{n4} e_i \Delta LnEX_{t-i}^+ + \sum_{i=0}^{n5} f_i \Delta LnEX_{t-i}^-$$

$$+ \theta_0 LnM_{t-1} + \theta_1 LnY_{t-1} + \theta_2 \pi_{t-1} + \theta_3 LnEX_{t-1}^+ + \theta_4 LnEX_{t-1}^- + \xi_t \qquad (3)$$

Equation (3) now allows us to test whether exchange rate changes have asymmetric or symmetric effects on the demand for money in the short run as well as in the long run. Shin et al. (2014) justify applying Pesaran *et al.*'s (2001) bounds testing approach to (3) in order to test the symmetry hypothesis. Equation (3) is labeled as a Nonlinear ARDL model in which nonlinearity is basically introduced through partial sums.[5]

3. The Results

In this section, we first try to estimate the linear ARDL model outlined by equation (2). Quarterly data from Iran is used for estimation purposes. Following others, we impose a

[4] For other applications of this approach see Bahmani-Oskooee et al. (2005), Bahmani-Oskooee and Hegerty (2007), Halicioglu, F., (2007), Narayan et al. (2007), Tang (2007), Mohammadi et al. (2008), Wong and Tang (2008), De Vita and Kyaw (2008), Payne (2008), Chen and Chen (2012), and Wong (2013).

[5] For another application of partial sum see Apergis and Miller (2006).

maximum of four lags on each first-differenced variable and use Akaike's Information Criterion (AIC) to select the optimum number of lags. The results from the optimum model are reported in Table 1.[6] Due to the volume of the results, they are reported in three panels. Panel A reports the short-run coefficient estimates. Clearly, a majority of these coefficients are significant, except the coefficients obtained for inflation rate and the exchange rate. Turning to the long-run normalized coefficients in Panel B, it appears that only income carries a positive and significant coefficient, supporting the transaction demand for money. But this long-run relation between money demand and income will be meaningful only if cointegration is established. From Panel C, it is clear that the variable addition F test for joint significance of lagged level variables at optimum lags is much less than its critical value of 3.52, implying the lack of cointegration. However, following Bahmani-Oskooee and Tanku (2008) we use normalized long-run coefficients from Panel B and equation (1) to generate the error term, which is normally referred to as error correction term *ECM*. We then replace the lagged level variables in (2) by ECM_{t-1} and estimate this new specification at the optimum lags. A significantly negative coefficient obtained for ECM_{t-1} will support cointegration or convergence toward the long run. This is indeed the case from Panel C.

A few other diagnostic statistics are also reported in Panel C. To make sure that the residuals are autocorrelation free, the Lagrange Multiplier (LM) statistic is reported, which is distributed as χ^2 with four degrees of freedom. Given its critical value of 9.48, reported at the bottom of Table 1, clearly the reported LM statistic is insignificant, supporting autocorrelation free residuals. Ramsey's RESET statistic is also reported for judging misspecification. This statistic also has a χ^2 distribution, but only with one degree of freedom. Given its critical value of 3.84 at the 5% level of significance, clearly our calculated statistic is insignificant, supporting the specification of the optimum model. We also report a test for normality of the residuals, which is based on skewness and kurtosis of the residuals. It has a χ^2 distribution with two degrees of freedom and, as can be seen, the reported statistic is insignificant, supporting the normality assumption. To test for the stability of short-run and long-run coefficients, following others (e.g., Pesaran *et al*. 2001, Bahmani-Oskooee and Bohl 2000) we apply CUSUM and CUSUMSQ tests to the residuals of the optimum model. As can be seen, all coefficients are stable by both criteria. Finally, adjusted R^2 is also reported, which shows a good fit.

[6] As a preliminary exercise, we applied the ADF test to each variable and tested for unit roots. The results, available upon request, revealed that all variables are I(1), except the inflation rate which was I(0).

Concentrating on the effects of exchange rate changes on the demand for money, based on the results so far, we may conclude that it has no significant effect during the study period. As mentioned before, this conclusion is based on assuming that the adjustment of variables is in linear fashion. Will the results change if we shift to nonlinear ARDL approach and estimate equation (3) by taking the same steps of imposing a maximum of four lags on each first differenced variable and using the AIC criterion? The results are reported in Table 2. From the short-run results, reported in Panel A, we gather that the variable representing negative partial sum, i.e., $\Delta LnEX^-$, which represents dollar depreciation, carries a significantly positive coefficient. The fact that $\Delta LnEX^-$ carries a significant coefficient but $\Delta Ln\ EX^+$ does not, supports asymmetry effects of any exchange rate changes on the demand for money in the short run. Does this last into the long run?

Long run coefficient estimates are reported in Panel B. Concentrating on the effects of exchange rate changes, it is clear that at the 5% significant level the negative partial sum, i.e., $LnEX^-$, carries a highly significant coefficient but the positive partial sum, i.e., $Ln\ EX^+$, does not. This is the same as the short-run results and it implies that when the dollar depreciates, Iranians hold less of their domestic currency.[7] If we consider the 10% level of significance, then both coefficients are significant with opposite signs, supporting, again, the asymmetric effects. The negative coefficient of positive partial sum, $Ln\ EX^+$, implies that when the dollar appreciates, Iranians hold less of their own currency. The positive coefficient of negative partial sum, $Ln\ EX^-$, also implies that when the dollar depreciates, Iranians hold less of their own currency. The combined implication of the results is that exchange rate changes have their effects on the demand for money in Iran through changes in expectations and not through the wealth effect. When the dollar appreciates, due to the expectation of further appreciation, Iranian will hold more dollars and less rials. When it depreciates, since there is always expectation of dollar appreciation, they will hold more dollars and less rial. Historically, due to inflation in Iran dollar has always appreciated. Any dollar depreciation has been short lived and there is a belief that like housing market in Iran, the trend of the dollar is always upward.[8] Therefore, we conclude that the exchange rate changes have an asymmetric effect on the demand for money in the long run.

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[7] Note that by way of construction, the negative partial sum can only decline. As it does, because of positive slope coefficient, so does the demand for rial.

[8] The 70 rials per dollar rate that existed before 1979, stands at more than 35000 rials per dollar as of revising this paper.

Turning to the other variables, it is clear that income elasticity is highly significant and as expected, it is positive. Inflation rate carries its expectedly negative and significant coefficient also. Like before, these long-run results will be meaningful only if we establish cointegration among variables. From Panel C, of Table 2, it is clear that cointegration is supported by at least ECM_{t-1} coefficient. Furthermore, all diagnostic statistics are less than their critical values, supporting autocorrelation free residuals, a correctly specified nonlinear ARDL model, normally distributed errors, and stable coefficients.[9]

4. Summary and Conclusion

The origin of including the exchange rate in the specification of the money demand goes back to 1963 and Robert Mundell, who conjectured that the demand for money could depend on the exchange rate in addition to income and interest rate. Subsequent studies since then have tried to provide a clear explanation along with empirical support for Mundell's conjecture. In this paper, we demonstrate that failure to find a significant relationship between the exchange rate and the demand for money could stem from assuming a linear dynamic adjustment process of variables toward their long run equilibrium variables. Once nonlinearity is introduced, using the partial sum concept, exchange rate movement could have a significant effect on the demand for money. We demonstrate this using Iranian data and by estimating the demand for money in Iran. Separating the movements in the value of the dollar into positive partial sum (appreciation) and negative partial sum (depreciation), helps us to introduce nonlinearity into the adjustment mechanism, we then find that dollar appreciation and dollar depreciation have an asymmetric effect on the demand for money in Iran.

Prior to the Iranian revolution, there was free capital mobility in Iran and, as a matter of fact, the Iranian rial was convertible and was one of the 16 currencies included in SDR (Special Drawing Rights) of the IMF. Iranians were allowed to hold assets abroad. Therefore, it is no surprise why when Bahmani-Oskooee (1996) used data over the period 1959-90, which was dominated by the pre-revolutionary period, support was found for the wealth effect of dollar appreciation. Due to sanctions, most Iranians cannot hold assets abroad, so the wealth effect is weakened. However, the expectation effect is very strong. Many Iranians in large cities now speculate by buying and selling dollars. This is borne out of our results in this paper, which used data only from the post-

[9] For sensitivity analysis, we carried out the entire analysis by imposing a maximum of eight lags on each first-differenced variable. There was no significant change in the results and conclusion. More precisely, income elasticity was 0.7915 with a t-ratio of 25.26; the coefficient obtained for inflation rate was -0.6667 with a t-ratio of

revolutionary period. As the dollar appreciates, Iranians expect further appreciation and therefore, hold more dollars and less of domestic currency.[10] The same is true even when the dollar depreciates, hence asymmetric effects.

Appendix

All data are quarterly and are collected from the following source.

a. The International Financial Statistics of the IMF.

b. Bahmani-Oskooee (2005).

Variables:

Real Money (M2): Nominal monetary aggregate (M1) is added to quazi money to get nominal M2. Nominal M2 is then deflated by CPI (the only available price index in Iran) to obtain real money. All data come from source a.

Real Income (Y): This variable is constructed by dividing the nominal GDP by the CPI. Quarterly nominal GDP was not available for Iran. Therefore, we generated the series following the method in Bahmani-Oskooee (1998) using annual data. All annual data came from source a.

Inflation Rate (Π) = This variable is defined as Ln P – Ln P_{t-1} , where P is measured by the CPI index. All CPI data come from source a.

Exchange Rate (EX) = This variable is defined as the number of Iranian rials per U.S. dollar, units of domestic currency per U.S. dollar. In line with Bahmani-Oskooee (1996), the black market exchange rate is used. Of course, since 2001 the official and the black market rates have been unified. The data come from sources (a) and (b).

-1.81, the coefficient of positive partial sum, i.e., $LnEX^+$, was -0.07389 with a t-ratio of -1.52, and finally the coefficient of negative partial sum, *Ln EX-*, was 0.4342 with a t-ratio of 2.91.

[10] Recent nuclear deals with the United Nations and the west and easing of sanctions has created optimism in Iran and has helped the dollar to come down from a high of 38,000 rials to almost 30,000.Part of that drop is due to the

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dollar selling by speculators. However, due to drop in oil prices and therefore, oil revenues which has reduced the supply of dollar to foreign exchange market, the rate is back to around 35000 rials per dollar.

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Table 1: Full-Information Estimate of Linear ARDL Equation (3).

Panel A: Short-Run Coefficient Estimates

	Lag Order				
	0	1	2	3	4
Δ Ln M	-	0.3316 (3.56)	-0.3904 (4.42)	0.3622 (3.90)	
Δ Ln Y	0.9805 (62.8)	-0.3611 (3.99)	0.3925 (4.42)	-0.3532 (3.82)	
Δπ	-0.0088 (0.49)				
Δ Ln EX	0.0012 (0.29)				

Panel B: Long-Run Coefficient Estimates

Constant	*Ln Y*	*π*	*Ln EX*
-4.6707 (2.01)	0.7705 (4.62)	-3.3997 (1.04)	-0.0631 (0.66)

Panel C: Diagnostic Statistics

F	ECM_{t-1}	*LM*	*RESET*	*Normality*	CUS (CUS^2)	Adj. R^2
1.11	-0.0115 (2.14)	7.33	0.1178	2.1417	Stable	0.99

Notes: a. The number inside the parentheses are the absolute value of t-ratios.
b. The upper bound critical value of the F statistic at the usual 5% level of significance is 3.52. This comes from Pesaran *et al.* (2001, Table CI-Case III, p. 300).
c. LM is the Lagrange multiplier test for serial correlation. It has a χ^2 distribution with four degrees of freedom. The critical value at the 5% level of significance is 9.48.
d. RESET is Ramsey's specification test. It has a χ^2 distribution with only one degree of freedom. The critical value at the 5% level of significance is 3.84.
e. The normality test is based on a test of skewness and kurtosis of residuals. It has a χ^2 distribution with only two degrees of freedom. The critical value at the 5% level of significance is 5.99.

Table 2: Full-Information Estimate of Nonlinear ARDL Equation (5).

Panel A: Short-Run Coefficient Estimates

	Lag Order				
	0	1	2	3	4
Δ Ln M	-	0.3390 (3.86)	-0.3320 (3.82)	0.3613 (4.32)	
Δ Ln Y	0.9710 (66.7)	-0.3682 (4.25)	0.3331 (3.82)	-0.3567 (4.28)	
Δπ	-0.0111 (0.67)				
Δ Ln EX$^+$	- 0.0156 (1.30)				
Δ Ln EX$^-$	0.0668 (2.73)				

Panel B: Long-Run Coefficient Estimates

Constant	*Ln Y*	*π*	*Ln EX*$^+$	*Ln EX*$^-$
-5.1661 (15.1)	0.8063 (37.9)	-0.6155 (1.96)	-0.0669 (1.85)	0.5035 (4.77)

Panel C: Diagnostic Statistics

F	*ECM*$_{t-1}$	*LM*	*RESET*	*Normality*	*CUSUM (CUSSUM*^{2}*)*	*Adj.* R^2
1.52	-0.0794 (2.81)	6.21	0.0303	2.4223	Stable	0.99

Notes: a. The number inside the parentheses are the absolute value of t-ratios.
b. The upper bound critical value of the F statistic at the usual 5% level of significance is 3.52. This comes from Pesaran *et al.* (2001, Table CI-Case III, p. 300).
c. LM is the Lagrange multiplier test for serial correlation. It has a χ^2 distribution with four degrees of freedom. The critical value at the 5% level of significance is 9.48.
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