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# Oil prices and exchange rates in oil-exporting countries: evidence from TAR and M-TAR models

**Hassan Mohammadi · Mohammad R. Jahan-Parvar**

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**Abstract** The paper studies the long-run relation and short-run dynamics between real oil prices and real exchange rates in a sample of 13 oil-exporting countries. The purpose of the study is to examine the possibility of Dutch disease in these countries. Tests of cointegration using threshold and momentum-threshold autoregressive (TAR and M-TAR) models suggest the possibility of the disease in 3-out-of 13 countries—Bolivia, Mexico and Norway. For these countries, we also find that (a) oil prices have a long-run effect on the exchange rates; and (b) exchange rates adjust faster to positive deviations from the equilibrium; and (c) there is no evidence of short-run causality between real exchange rates and real oil prices in either direction. Over all, these findings suggest a weak link between oil prices and real exchange rates and thus limited evidence in favor of the Dutch disease.

**Keywords** Asymmetry · Cointegration · Dutch Disease · Error Correction · Oil Prices · Real Exchange Rates · Threshold and Momentum Threshold Autoregressive Models

JEL Classification · C32 · C52 · F31 · F37 · F47

## 1 Introduction

The past few years has witnessed a revival of interest in the effect of high oil prices on the economies of oil-exporting-countries. Most literature on the subject falls into

H. Mohammadi (✉)
Department of Economics, Illinois State University, Campus Box 4200, Normal, IL 61790-4200, USA
e-mail: hmohamma@ilstu.edu

M. R. Jahan-Parvar
Department of Economics, East Carolina University, A-426 Brewster Building, Greenville, NC 27858, USA
e-mail: jahanparvarm@ecu.edu

two broadly related categories. One is the ongoing literature on the "curse of natural resources" (Sachs and Warner 2001), which postulates that resource-rich countries tend to under-perform their resource-poor counterparts in terms of economic development and poverty reduction. The other is the literature on "Dutch disease" (Buiter and Purvis 1982; Corden and Neary 1982; Corden 1984) which is primarily concerned with potential macroeconomic effects of the curse, such as the appreciation of the real exchange rate, loss of competitiveness in export oriented sectors, decline in the share of manufacturing industry and a boom in non-tradables in afflicted economies.[1] As it appears, the appreciation of the real exchange rate serves as the necessary condition for "Dutch disease".[2]

Beginning with oil shocks of the early 1970s, a large and growing body of studies has focused on the experience of oil-exporting countries. Early studies include Buiter and Purvis (1982), Bruno and Sachs (1982), Edwards and Aoki (1983), Neary and Van Wijnbergen (1986), and Warr (1986). More recent research includes Fardmanesh (1991), Hutchinson (1994), Brunstad and Dyrstad (1997), Bjorland (1998), Koranchelian (2005), Issa et al. (2008), Zalduendo (2006), Kalcheva and Oomes (2007), Jahan-Parvar and Mohammadi (2009, 2010) and Korhonen and Juurikkala (2009).

As it appears, evidence based on traditional tests of cointegration generally supports the Dutch disease hypothesis. For example, Koranchelian (2005) finds strong support for the effect of oil prices on real exchange rates in Algeria. Zalduendo (2006) reports similar results for Venezuela. Issa et al. (2008) find that higher energy prices resulted in an appreciation of Canadian dollar since 1993 when Canada became a net energy exporter. Kalcheva and Oomes (2007) report similar results for Russia with long-run elasticity of real exchange rates with respect to real oil prices around 0.5. Korhonen and Juurikkala (2009) reach similar conclusions using a pooled mean group estimator applied to a panel data set of nine OPEC countries. More recently, Jahan-Parvar and Mohammadi (2010) examine the possibility of Dutch disease in 14 oil-exporting countries by applying the autoregressive distributed lag model of Pesaran et al. (2001). Their findings also

[1] "Dutch disease," refers broadly to the harmful consequences of exogenous increases in a country's income and wealth due to discovery of natural resources, inflow of foreign currency due to a surge in natural resource prices, foreign assistance, and foreign direct investment. These developments, although positive, may appreciate the country's real exchange rate, and as a result crowd out the traditional exporting sector. For example, a rise in oil prices increases the flow of foreign exchange in the oil exporting country, and raises its income. If the foreign exchange is spent entirely on imports, it would have no direct impact on supply of money or demand for domestically produced goods. However, it if is converted into domestic currency and spent on non-traded goods, it will result in real exchange rate appreciation irrespective of the choice of exchange rate regime. With flexible exchange rates, the appreciation occurs through a rise in nominal value of domestic currency while with fixed exchange rates it occurs through a rise in domestic prices. The appreciation of real exchange rates weakens the competitiveness of the country's exports, crowds out the traditional export sector, and reallocates resources from the exports sector to the non-tradable sector (Ebrahim-zadeh 2003).

[2] The effect of oil prices on real exchange rates may be considered as a part of a broader literature on exchange rate determination and tests of purchasing power parity. According to this literature, the real exchange rate may be interpreted as a measure of deviations from purchasing power parity (PPP). In such framework, a stationary real exchange rate is consistent with the PPP theory. Some recent contributions to this literature include Taylor et al. (2001), Taylor and Sarno (2004), Taylor and Taylor (2004), and Lopez (2008).

support the existence of long-run relations between oil prices and exchange rates. However, estimates of long-run elasticities are meaningful and significant only for four countries of Angola, Indonesia, Nigeria and Russia, and range between 1.813 for Angola to 0.355 for Indonesia. These results cast a shadow of doubt on the extent that real exchange rates respond to oil price shocks.[3]

A critical assumption underlying previous tests of cointegration is symmetric adjustments of exchange rates in response to positive and negative deviations from equilibrium. However, several factors may contribute to asymmetric responses in real exchange rates to oil prices in oil-exporting countries. As Enders and Dibooglu (2001) point out, official interventions may result in asymmetric adjustments in exchange rates. Under a managed float, monetary authorities might be more willing to tolerate a currency appreciation than depreciation. Also, the mere existence of currency bands may mitigate exchange rate movements unless the level of the band is altered. Similarly, stickiness of nominal prices, especially in the downward direction, provides additional justification for asymmetric adjustments in real exchange rates. Also, as Ewing et al. (2006) point out, asymmetries through heterogeneous expectations about nominal exchange rates by risk-averse market participants, high transactions costs which inhibit or slow the adjustment process; economic and political shocks which may affect oil prices and exchange rates differently; and institutional factors such as OPEC pricing and quota mechanisms may prevent exchange rates from moving in tandem with real oil prices.

Putting together, these factors provide sufficient justifications for the possibility of asymmetries in the dynamics of oil price—exchange rate relationship. In fact, a number of studies (Taylor and Taylor 2004; Taylor et al. 2001) have explored that possibility using data from industrial economies. Given asymmetric adjustments, dynamic relations implicit in tests of the Dutch disease hypothesis based on traditional methods of Engle and Granger (1987), Johansen (1988) and Pesaran et al. (2001) are mis-specified. Furthermore, traditional tests of cointegration will have low power (see for example, Pippenger and Goering 1993; Enders and Granger 1998; Enders and Siklos 2001). The proper approach is to allow for asymmetric response by applying threshold autoregressive (TAR) and momentum threshold autoregressive (MTAR) models developed by Enders and Granger (1998), Enders and Siklos (2001) and Enders (2001).[4]

This paper re-examines the existence of threshold cointegration between real oil prices and real exchange rates in a sample of 13 oil-exporting countries. We show that allowing for threshold adjustments yield results that are much less supportive of the Dutch disease hypothesis. Thus, previous empirical support may reflect a model mis-specification.

[3] A number of other studies have examined the effect of other commodity prices on equilibrium exchange rates of commodity-exporting countries. In particular, Cashin et al. (2004) find strong relationship between real commodity prices and real exchange rates in about a third of 58 commodity-exporting countries in their sample. Chen and Rogoff (2003) report a strong relationship between US dollar prices for commodity exports and exchange rates in three major commodity exporting OECD members—Australia, Canada and New Zealand.

[4] Recent applications of this method are Enders and Granger (1998), Enders and Dibooglu (2001), Bajo-Rubio et al. (2004), and Payne and Mohammadi (2006).

The paper proceeds as follows: Section 2 introduces the empirical model and discusses the contribution of threshold autoregressive models to the analysis. Section 3 describes data sources and construction of the variables. Section 4 discusses the empirical findings; and Section 5 concludes.

## 2 Empirical model

Conventional residual-based tests of cointegration (Engle and Granger 1987) examine the validity of Dutch disease hypothesis by estimating the following model,

$$e_t = \alpha_0 + \alpha_1 q_t + \varepsilon_t \tag{1}$$

$$\Delta\varepsilon_t = \rho\varepsilon_{t-1} + \sum_{i=1}^{k} \beta_i \Delta\varepsilon_{t-i} + v_t \tag{2}$$

where $e_t$ is the log of real exchange rates and measures the relative price of domestic goods with respect to a basket of foreign goods such that a rise in$e_t$ corresponds to a real exchange rate appreciation; $q_t$ is the log of real oil prices; $\varepsilon_t$ is a stochastic innovation term; and$\alpha_1$is the elasticity of real exchange rates with respect to real oil prices.[5]

Equation 1 represents the long-run relationship between oil prices and real exchange rates.[6] Support for the hypothesis requires that $\alpha_1>0$, and $\varepsilon_t$ follow a stationary process. The latter condition is satisfied if the null hypothesis of unit roots ($\rho=0$) in the augmented Dickey and Fuller (1979) model in Eq. 2 is rejected in favor of the alternative that $\rho<0$.

An implicit assumption in tests of unit roots in (2) is the linear and symmetric adjustments of the exchange rates to positive and negative deviations from the equilibrium. In contrast, several studies have documented non-linear adjustments in real exchange rates (see for example, Taylor et al. 2001), implying that Eq. 2 is misspecified. An alternative which allows for asymmetric adjustments is the TAR and M-TAR models as in Enders and Granger (1998), Enders (2001), and Enders and Siklos (2001). That requires estimating the following modified ADF model:

$$\Delta\varepsilon_t = \rho^+ M_t \varepsilon_{t-1} + \rho^- (1 - M_t)\varepsilon_{t-1} + \sum_{i=1}^{k} \beta_i \Delta\varepsilon_{t-i} + u_t \tag{3}$$

where $M$ is the Heaviside indicator; and parameters $\rho^+$and $\rho^-$ allow for asymmetric autoregressive decay. The criteria for the Heaviside indicator differ between TAR

[5] Estimation of Eq. 1 requires assumptions regarding the stochastic properties of real oil prices. Given the global nature of oil markets and potential stickiness in domestic prices, it is reasonable to assume that real oil prices are exogenous with respect to real exchange rates.

[6] According to Taylor et al. (2001), the real exchange rate may be interpreted as a measure of the deviation from purchasing power parity. As such, Eq. 1 explains such deviations in terms of real oil prices.

and MTAR models. In the TAR model, it is set according to the previous value of the error term relative to an optimum threshold, $\tau$:

$$\begin{cases} M_t = 1 & \text{if } \varepsilon_{t-1} \geq \tau \\ \quad\; = 0 & \text{if } \varepsilon_{t-1} < \tau \end{cases} \tag{4}$$

In contrast, in the M-TAR model, it is set according to the change in the previous value of the error term relative to an optimum threshold:

$$\begin{cases} M_t = 1 & \text{if } \Delta\varepsilon_{t-1} \geq \tau \\ \quad\; = 0 & \text{if } \Delta\varepsilon_{t-1} < \tau. \end{cases} \tag{5}$$

where $\tau$ is set endogenously using Chan (1993) method. As is clear from (4) and (5), the degree of autoregressive decay in the models depends on the state of the equilibrium error.

The test procedure consists of three steps: The first step involves estimating the TAR and M-TAR models and testing for cointegration. The null hypothesis of no-cointegration ($H_o$:$\rho^+$=$\rho^-$=0) is examined by comparing the actual values of the test statistics $\Phi_\mu$ with their corresponding critical values. If the null hypothesis of no-cointegration is rejected, then one proceeds to the second step, which involves testing the null hypothesis of symmetry ($H_o$:$\rho^+$=$\rho^-$). If the null hypothesis of symmetry is rejected and $|\rho^+|>|\rho^-|$, then there is evidence of relatively more decay in response to positive deviations from equilibrium.

The third step involves the estimation of the asymmetric error-correction model and tests of long-run and short-run causality between real oil prices ($q_t$) and real exchange rates ($e_t$). The asymmetric error-correction model consists of two equations as follow:

$$\Delta e_t = \delta_1^+ M_t \varepsilon_{t-1} + \delta_1^- (1 - M_t)\varepsilon_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta e_{t-i} + \sum_{i=1}^{k} \varphi_i \Delta q_{t-i} + u_t \tag{6}$$

$$\Delta q_t = \delta_2^+ M_t \varepsilon_{t-1} + \delta_2^- (1 - M_t)\varepsilon_{t-1} + \sum_{i=1}^{k} \gamma_i^* \Delta e_{t-i} + \sum_{i=1}^{k} \varphi_i^* \Delta q_{t-i} + u_t^* \tag{7}$$

Equation 6 represents the error-correction model for the real exchange rate. Here the change in real exchange rate is modeled in terms of positive ($M_t\varepsilon_{t-1}$) and negative ($(1-M_t)\varepsilon_{t-1}$) lagged equilibrium errors as well as past changes in the real exchange rate and in real oil prices. Real oil prices Granger-cause real exchange rates in the long-run if the coefficients of lagged equilibrium errors ($\delta_1^+$ and $\delta_1^-$) are jointly significant, and in the short-run if the coefficients of the past oil price changes ($\varphi_i (i=1,...,k)$) are jointly significant.

Similarly, error-correction Eq. 7 can be used to examine long-run and short-run causality from real exchange rates to real oil prices. Here the change in real oil prices is modeled in terms of positive and negative lagged equilibrium errors as well as past changes in real oil prices and in the real exchange rate. Long-run causality from real exchange rates to real oil prices is examined by testing the joint significance of $\delta_2^+$ and $\delta_2^-$ parameter estimates. The null-hypothesis of no long-run causality is rejected if the two parameter estimates are jointly insignificant. Similarly, short-run causality from changes in real exchange rates to changes in real oil prices in examined by testing the joint significance of $\gamma_i^*(i=1,...,k)$parameter estimates. The null hypothesis of no short-run causality is rejected if $\gamma_i^*(i=1,...,k)$are jointly insignificant.

The analysis so far assumes non-stationary yet cointegrated series. In the absence of cointegration, we proceed to tests of short-run causality using a vector autoregressive (VAR) model with first differenced data (Eqs. 8 and 9):

$$\Delta e_t = \sum_{i=1}^{k} \gamma_i \Delta e_{t-i} + \sum_{i=1}^{k} \varphi_i \Delta q_{t-i} + u_t \tag{8}$$

$$\Delta q_t = \sum_{i=1}^{k} \gamma_i^* \Delta e_{t-i} + \sum_{i=1}^{k} \varphi_i^* \Delta q_{t-i} + u_t^* \tag{9}$$

where the variables and parameters are defined as before. As for stationary variables, we proceed to tests of short-run causality using the corresponding VAR in levels.

## 3 Data

Monthly data on nominal exchange rates, consumer price indices (CPI) and nominal price of oil are collected from the International Financial Statistics of the IMF data bank. We measure real exchange rates by price of a basket of domestically produced goods relative to a basket of goods produced in the U.S.[7] Thus, an increase in the real exchange rate implies real appreciation for the home country. We measure real oil prices by nominal price of West Texas Intermediate (henceforth WTI) deflated by the U.S. consumer price index.[8] Natural logs of real oil prices and real exchange rates are used in the investigation. As reported in column 2 of Table 1, data availability varies across the 13 countries.[9] The longest sample period begins in 1970:1 and ends in 2010:1 and covers seven of the countries (Colombia, Gabon, Indonesia, Mexico, Nigeria, Norway and Venezuela). These countries are listed in panel A of Table 1. As for the remaining six countries, the starting period varies across countries: 1974:1 for Algeria, 1995:11 for Angola, 1986:1 for Bolivia, 1973:1 for Kuwait, 1992:9 for Russia, and 1980:2 for Saudi Arabia. These countries are listed in panel B of the Table.

[7] More specifically, let $P$, $P^*$ and $E$ represent domestic price, foreign price, and the nominal exchange rate (price of a unit of foreign currency in terms of domestic currency), respectively. Then the real exchange rate is $P/E\times P^*$.

[8] As a check of robustness, we also experimented with two alternative oil prices, the North Sea Brent and the Dubai Fateh spot prices. The results are robust to this choice.

[9] Bahrain was included in the original sample. However, due to data limitations, it was excluded in these analyses.

**Table 1** Traditional tests of unit roots

| | | Level | | | First difference | | |
|---|---|---|---|---|---|---|---|
| Variable | Sample Period | ADF (Constant) | ADF (Trend) | PP (Constant) | ADF (Constant) | ADF (Trend) | PP (Constant) |
| Panel A | | | | | | | |
| e_Colombia | 1970:01–2010:01 | −1.889 [12] | −1.999 [12] | −1.349 [0] | −4.159* [12] | −4.182* [12] | −14.185* [0] |
| e_Gabon | 1970:01–2010:01 | −0.684 [0] | −0.199 [0] | −1.281 [0] | −20.934* [0] | −21.092* [0] | −22.306* [0] |
| e_Indonesia | 1970:01–2010:01 | −0.947 [0] | 0.938 [0] | −1.407 [0] | −12.925* [1] | −12.899* [1] | −19.996* [0] |
| e_Mexico | 1970:01–2010:01 | −2.607 [4] | −3.120 [10] | −3.104 [0] | −5.853* [9] | −5.849* [9] | −20.507* [0] |
| e_Nigeria | 1970:01–2010:01 | −1.439 [0] | −2.045 [0] | −1.644 [0] | −19.718* [0] | −19.741* [0] | −21.011* [0] |
| e_Norway | 1970:01–2010:01 | −2.562 [0] | −2.563 [0] | −2.607 [0] | −20.541* [0] | −20.522* [0] | −20.944* [0] |
| e_Venezuela | 1970:01–2010:01 | −2.411 [2] | −2.495 [2] | −2.643 [0] | −17.788* [1] | −14.771* [2] | −24.319* [0] |
| q | 1970:01–2010:01 | −2.677 [1] | −2.414 [12] | −1.713 [0] | −18.162* [0] | −18.141* [0] | −17.589* [0] |
| Panel B | | | | | | | |
| e_Algeria | 1974:01–2010:01 | −0.739 [1] | −2.355 [12] | −0.579 [0] | −4.929* [2] | −4.909* [12] | −18.477* [0] |
| e_Angola | 1995:11–2010:01 | −2.046 [0] | −2.291 [0] | −2.429 [0] | −10.531* [0] | −6.635* [5] | −10.790* [0] |
| e_Bolivia | 1986:01–2010:01 | −1.176 [0] | −1.877 [0] | −3.000 [0] | −10.888* [0] | −10.842* [0] | −13.089* [0] |
| e_Kuwait | 1973:01–2010:01 | −0.033 [0] | −1.838 [12] | −2.571 [0] | −5.120* [0] | −5.268* [0] | −20.605* [0] |
| e_Russia | 1992:09–2010:01 | −2.189 [9] | −2.310 [9] | −1.713 [0] | −4.654* [8] | −4.493* [8] | −10.449* [0] |
| e_Saudi Arabia | 1980:02–2010:01 | −4.038* [1] | −2.736 [1] | −4.255 [0] | −14.195* [0] | −14.805* [0] | −14.676* [0] |

*e* and *q* stand for real exchange rate and real oil prices respectively

SIC based lag-lengths are in brackets

Significance at 5% level is represented by *

## 4 Empirical results

We begin with tests of unit roots in real exchange rates and real oil prices using traditional augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. Table 1 reports the results of these tests. All test statistics overwhelmingly fail to reject the null hypothesis of unit roots in levels but reject the hypothesis with first-differenced data. The only exception is the real exchange rate for Saudi Arabia, where the null of unit roots in level is rejected by the ADF without trend.

A potential shortcoming of the ADF and PP tests is that they do not account for structural breaks in the time series under investigation. However, the importance of allowing for structural breaks in unit root tests is now well documented. This is a potentially serious shortcoming given that several of the countries in the sample faced major currency crisis during the sample period. Thus, failure to reject unit roots in real exchange rates might be due to breaks in the series in response to currency crisis.[10] To address this issue, we re-examine the robustness of the previous test results using three alternative unit root tests with endogenous structural breaks, namely Lee and Strazicich (2003) (henceforth, LS), Zivot and Andrews (1992) (henceforth, ZA), and Perron (1997) (henceforth, P). The results of these tests are reported in Table 2. For the seven countries with the longest sample period, reported in panel A, the results fail to reject the null hypothesis of unit roots in real exchange rates, and are consistent with findings from the traditional tests. The only exceptions are Mexico and Venezuela, where the LS statistic rejects the null hypothesis. As for the remaining six countries with shorter sample periods (reported in panel B), the null hypothesis of unit roots fails to be rejected across all three tests for Algeria and Russia; and across ZA and LS tests for Bolivia and Saudi Arabia. The only exception is Angola, where the null hypothesis is rejected across all three tests.

Given this evidence, we proceed by treating all real exchange rates as non-stationary in level and stationary in first-differences, and proceed to tests of cointegration. The only exception is the exchange rate for Angola which we treat as stationary, and exclude it from tests of cointegration.

Tables 3 and 4 report the results of tests of cointegration between exchange rates and oil prices under TAR and MTAR adjustments, respectively. We begin with the results of the TAR model in Table 3. Three patterns emerge: first, among the seven countries in panel A, the calculated $\Phi_\mu$- statistic for the null hypothesis of no-cointegration ($H_o$:$\rho^+$=$\rho^-$=0) exceeds its 5% critical value of 6.32 for Norway, and its 10% critical value of 5.21 for Mexico. Similarly, among the five countries in panel B, the null hypothesis of no-cointegration is rejected for Bolivia at 5% significance level. Thus, evidence in favor of long-run relations between oil prices and exchange rates under asymmetric TAR adjustments is limited to only three countries. Second, among these three countries, we also reject the null hypothesis of symmetric adjustments in equilibrium errors ($H_o$:$\rho^+$=$\rho^-$) for Norway and Bolivia. Third, for these two countries, the estimates of $\rho^+$ is statistically larger than zero in absolute value and exceed the value of $\rho^-$, suggesting that equilibrium errors adjust faster in response to positive deviations from equilibrium.

[10] We thank two referees for pointing out the possibility of currency crisis, the existence of endogenous breaks in the real exchange rates, and thus tests of unit roots with endogenous breaks.

**Table 2** Tests of unit roots with endogenous breaks

| Variable | Sample Period | Perron | Zivot-Andrews | Lee-Strazicich |
|---|---|---|---|---|
| Panel A | | | | |
| e_Colombia | 1970:01–2010:01 | −2.085 (1983:03) | −3.169 (1983:06) | −1.538 (1994:03) |
| e_Gabon | 1970:01–2010:01 | −2.483 (2007:02) | −2.080 (1985:03) | −2.274 (1993:12) |
| e_Indonesia | 1970:01–2010:01 | −2.894 (2007:08) | −0.470 (2001:07) | −2.255 (1986:08) |
| e_Mexico | 1970:01–2010:01 | −5.436 (1981:12) | −4.705 (1982:02) | −3.990* (1982:01) |
| e_Nigeria | 1970:01–2010:01 | −3.955 (1986:04) | −3.769 (1986:06) | −2.016 (1986:09) |
| e_Norway | 1970:01–2010:01 | −3.547 (1996:12) | −3.228 (1992:09) | −2.027 (1997:03) |
| e_Venezuela | 1970:01–2010:01 | −4.943 (1986:10) | −4.792 (1984:02) | −3.664* (1984:01) |
| q | 1970:01–2010:01 | −3.721 (1985:10) | −3.543 (1985:12) | −2.399 (1986:01) |
| Panel B | | | | |
| e_Algeria | 1974:01–2010:01 | −4.271 (2007:08) | −2.958 (1990:05) | −2.598 (1991:08) |
| e_Angola | 1995:11–2010:01 | −9.109* (1999:03) | −7.669* (1999:05) | −4.761* (1999:04) |
| e_Bolivia | 1986:01–2010:01 | −6.052* (2007:08) | −3.662 (2006:03) | −3.365 (1989:05) |
| e_Kuwait | 1973:01–2010:01 | −8.861* (1978:11) | −8.806* (1979:01) | −2.658 (1978:02) |
| e_Russia | 1992:09–2010:01 | −4.593 (2007:08) | −4.184 (2007:06) | −1.935 (2007:09) |
| e_Saudi Arabia | 1980:02–2010:01 | −48.308* 2007:08 | −2.493 (2005:08) | −1.753 (1994:12) |

*e* and *q* stand for real exchange rate and real oil prices respectively

Significance at 5% level is represented by *

The 5% critical values are: −5.550 for Perron; −4.800 for Zivot-Andrews; and −3.566 for Lee-Strazicich tests

The findings from the M-TAR model, reported in Table 4, provides broadly similar conclusions with three exceptions: first, the support for cointegration is now limited to only two countries, Mexico and Bolivia at the 5% level of significance. Second, for these two countries, there is also support for asymmetric adjustments in equilibrium errors. And third, the estimates of $\rho^+$ exceeds those of $\rho^-$ in absolute values, thus reinforcing the TAR results that equilibrium errors adjust faster to positive deviations from equilibrium.

Finally, we report the results of tests of Granger non-causality between exchange rates and oil prices in Table 5. Panel A reports the estimates of asymmetric error-correction parameters along with $F$-statistics associated with tests of long-run and short-run causality for the three countries with cointegrated relations; Bolivia, Mexico and Norway. Panel B reports the $F$-statistics associated with short-run causality with first-differenced data for the nine countries with non-stationary variables and in the absence of cointegration. Finally, panel C reports tests of causality with level data for Angola, where the real exchange rate appeared stationary.

Our main findings are summarized as follow: First, evidence from panel A supports the existence of long-run causality from real oil prices to real exchange rates in each of the three countries with cointegrated relations. This is evident in the calculated values of $F(\delta_1^+=\delta_1^-=0)$ for the joint significance of $\delta_1^+$ and $\delta_1^-$ (in Eq. 6), which are significantly larger than their corresponding 5% critical values. In contrast, the calculated values of $F(\delta_2^+=\delta_2^-=0)$ for the joint significance of $\delta_2^+$ and

**Table 3** Tests of cointegration and asymmetric adjustments: TAR model

| Country | Sample | $\tau$ | $\rho^+$ | $\rho^-$ | $\Phi_\mu^*$ | t–stat | $Q_{LB}(12)$ |
|---|---|---|---|---|---|---|---|
| A. Countries with the long sample period | | | | | | | |
| Colombia (12) | 1970:01–2010:01 | 0.168 | −0.028 (0.013)* | −0.008 (0.009) | 2.465 | 1.221 | 1.189 [0.998] |
| Gabon (2) | 1970:01–2010:01 | 0.331 | −0.044 (0.015)* | −0.001 (0.009) | 4.510 | 2.559* | 5.646 [0.509] |
| Indonesia (12) | 1970:01–2010:01 | −1.268 | 0.001 (0.003) | 0.001 (0.004) | 0.068 | 0.251 | 1.058 [0.955] |
| Mexico (12) | 1970:01–2010:01 | 0.117 | −0.077 (0.028)* | −0.042 (0.017)* | 6.030** | 1.095 | 0.452 [0.995] |
| Nigeria (2) | 1970:01–2010:01 | −0.417 | −0.033 (0.014)* | −0.033 (0.015) | 4.768 | −0.049 | 0.062 [0.999] |
| Norway (2) | 1970:01–2010:01 | 0.139 | −0.066 (0.021)* | −0.017 (0.012) | 6.525* | 2.114* | 7.778 [0.704] |
| Venezuela (2) | 1970:01–2010:01 | 0.138 | −0.032 (0.023) | −0.035 (0.016)* | 3.204 | −0.119 | 6.987 [0.177] |
| B. Countries with varying sample periods | | | | | | | |
| Algeria (12) | 1974:01–2010:01 | 0.178 | −0.030 (0.019) | −0.044 (0.018)* | 3.980 | −0.559 | 0.181 [0.999] |
| Angola (2) | – | – | – | – | – | – | – |
| Bolivia (12) | 1986:01–2010:01 | 0.425 | −0.197 (0.054)* | 0.011 (0.031) | 7.301* | 3.667* | 1.498 [0.780] |
| Kuwait (2) | 1973:01–2010:01 | 0.181 | −0.035 (0.011)* | 0.006 (0.011) | 4.835 | 2.601 | 4.111 [0.309] |
| Russia (12) | 1992:09–2010:01 | 0.453 | −0.015 (0.027) | −0.047 (0.019)* | 2.833 | −0.985 | 0.728 [0.999] |
| Saudi Arabia (8) | 1980:02–2010:01 | 0.980 | 0.094 (0.033)* | −0.002 (0.015) | 3.870 | 2.536 | 0.099 [0.999] |

$\tau$ is optimum threshold; $\rho^+$ and $\rho^-$ are the adjustment coefficients; $\Phi_\mu^*$ is the test statistic for the null hypothesis that$\rho_1=\rho_2=0$. It's 5% and 10% critical value from Enders and Siklos (2001) are 6.32 and 5.21 respectively; *t* is the student test statistic for the null hypothesis of symmetry; and $Q_{LB}(12)$ is the Ljung-Box Q-Statistics for test of serial correlation

Numbers in parentheses under country names are auxiliary lag lengths

Numbers in parentheses next to coefficients are standard errors

Numbers in brackets are p-values

Significant at 5% and 10% levels are represented by * and ** respectively

**Table 4** Tests of cointegration and asymmetric adjustments: M-TAR model

| Country | Sample Period | $\tau$ | $\rho^{+}$ | $\rho^{-}$ | $\Phi_{\mu}^{*}$ | *t–stat* | $Q_{LB}(12)$ |
|---|---|---|---|---|---|---|---|
| A. Countries with the long sample period | | | | | | | |
| Colombia (12) | 1970:01–2010:01 | 0.017 | −0.038 (0.016)* | −0.008 (0.008) | 3.052 | 1.630 | 1.397 [0.998] |
| Gabon (2) | 1970:01–2010:01 | 0.032 | −0.046 (0.022)* | −0.008 (0.008) | 2.539 | 1.621 | 4.842 [0.980] |
| Indonesia (12) | 1970:01–2010:01 | −0.081 | −0.001 (0.070) | −0.001 (0.002) | 0.050 | 0.165 | 1.126 [0.999] |
| Mexico (12) | 1970:01–2010:01 | 0.021 | −0.189 (0.041)* | −0.029 (0.016)* | 12.375* | 3.685* | 0.894 [0.999] |
| Nigeria (2) | 1970:01–2010:01 | 0.051 | −0.092 (0.021)* | −0.016 (0.012) | 4.768 | −0.049 | 0.052 [0.999] |
| Norway (2) | 1970:01–2010:01 | −0.024 | −0.043 (0.027)* | −0.027 (0.011)* | 4.413 | 0.564 | 7.398 [0.704] |
| Venezuela (2) | 1970:01–2010:01 | −0.006 | −0.033 (0.029) | −0.034 (0.015)* | 3.197 | −0.024 | 6.989 [0.177] |
| B. Countries with varying sample periods | | | | | | | |
| Algeria (12) | 1974:01–2010:01 | 0.027 | −0.029 (0.031) | −0.039 (0.014)* | 3.865 | −0.294 | 0.149 [0.180] |
| Angola (2) | – | – | – | – | – | – | – |
| Bolivia (12) | 1986:01–2010:01 | 0.041 | −0.417 (0.072)* | −0.013 (0.029) | 17.360* | 5.787* | 1.998 [0.780] |
| Kuwait (2) | 1973:01–2010:01 | 0.011 | −0.032 (0.017)* | 0.009 (0.009) | 2.124 | 1.175 | 3.401 [0.909] |
| Russia (12) | 1992:09–2010:01 | −0.006 | −0.042 (0.037) | −0.036 (0.019)** | 2.351 | 0.167 | 0.849 [0.999] |
| Saudi Arabia (8) | 1980:02–2010:01 | −0.001 | 0.002 (0.028) | −0.022 (0.016) | 0.921 | −0.745 | 0.698 [0.999] |

See notes to Table 3

$\delta_2^{-}$ (in Eq. 7) are below their corresponding 5% critical values, suggesting the absence of long-run causality from real oil prices to real exchange rates. These findings imply the existence of uni-directional long-run causality from real oil prices to real exchange rates in countries in panel A. Second, the estimates of error-correction parameters associated with positive equilibrium errors are larger in absolute value than those associated with negative equilibrium errors. Thus, exchange rates may adjust faster to positive deviations from equilibrium. These findings reinforce the previous results based on tests of cointegration reported in Tables 3 and 4. Third, there is no evidence of short-run causality in either direction in the three countries in panel A. The calculated $F$-statistics associated with $F(\varphi_i=0)$ and $F(\varphi_i^{*}=0)$ are both below their 5% critical values. Fourth, there is also no evidence of short-run causality in either direction for the nine countries in panel B, as reflected in the small values for $F(\varphi_i=0)$ and $F(\varphi_i^{*}=0)$ for these countries. Thus, there is no support for the view that changes in oil prices affects real exchange rates in the short-run, or that changes in real exchange rates affect real oil prices. Perhaps the only exception is for Colombia where the latter hypothesis is rejected at the 5% significance level. Fifth, finally, as reported in panel C, there is no evidence of causality between levels of oil prices and exchange rates for Angola in either direction.

## 5 Conclusions

We study the relationship between real oil prices and real exchange rates in 13 oil exporting countries using threshold and momentum threshold autoregressive models. The purpose of the study is to examine the possibility of the Dutch disease in oil exporting countries.

**Table 5** Tests of long-run and short-run causality

| Country | $H_o$:$q \nRightarrow e$ | | | | $H_o$:$e \nRightarrow q$ | | | |
|---|---|---|---|---|---|---|---|---|
| | $\delta_1^+$ | $\delta_1^-$ | $F(\delta_1^+=\delta_1^-=0)$ | $F(\varphi_i=0)$ | $\delta_2^+$ | $\delta_2^-$ | $F(\delta_2^+=\delta_2^-=0)$ | $F(\varphi_i^*=0)$ |
| A. VECM with asymmetric adjustments | | | | | | | | |
| Bolivia | −0.206* (0.055) | 0.012 (0.034) | 7.425* [0.001] | 0.583 [0.856] | 0.035 (0.024) | −0.004 (0.015) | 1.139 [0.322] | 0.331 [0.983] |
| Mexico | −0.079* (0.026) | −0.044* (0.018) | 6.926* [0.001] | 0.919 [0.527] | 0.035 (0.045) | 0.012 (0.031) | 0.349 [0.705] | 0.294 [0.990] |
| Norway | −0044* (0.018) | −0.021* (0.012) | 4.619* [0.010] | 1.290 [0.276] | 0.017 (0.053) | 0.028 (0.036) | 0.348 [0.706] | 0.379 [0.685] |
| B. VAR with first-differenced variables | | | | | | | | |
| Algeria | | | | 1.125 [0.337] | | | | 0.764 [0.687] |
| Colombia | | | | 1.359 [0.182] | | | | 2.863* [0.001] |
| Gabon | | | | 2.809 [0.061] | | | | 1.743 [0.176] |
| Indonesia | | | | 1.412 [0.156] | | | | 1.280 [0.227] |
| Kuwait | | | | 1.585 [0.206] | | | | 0.561 [0.571] |
| Nigeria | | | | 0.261 [0.770] | | | | 2.663 [0.071] |
| Russia | | | | 1.231 [0.265] | | | | 0.442 [0.944] |
| Saudi Arabia | | | | 0.847 [0.561] | | | | 0.696 [0.695] |
| Venezuela | | | | 1.079 [0.341] | | | | 0.618 [0.539] |
| C. VAR with Level Variables | | | | | | | | |
| Angola | | | | 1.928 [0.149] | | | | 1.413 [0.246] |

Numbers in parentheses are estimated standard errors

Numbers in brackets are significance levels

Significant at 5% level is designated by '*'

$H_o$:$q \nRightarrow e$ refers to null of no-causality from oil prices to exchange rate; $\delta_1^+$ and $\delta_1^-$ are the corresponding equilibrium error adjustment parameters; $F(\delta_1^+=\delta_1^-=0)$ is the F-statistic associated long-run causality from oil prices to real exchange rates; $F(\varphi_i=0)$ is the F-statistic associated with short-run causality from oil prices to exchange rates

$H_o$:$e \nRightarrow q$ refers to null of no-causality from real exchange rates to real oil prices; $\delta_2^+$ and $\delta_2^-$ are the corresponding equilibrium error adjustment parameters; $F(\delta_1^+=\delta_1^-=0)$ is the F-statistic associated with long-run causality from real exchange rates to real oil prices; $F(\varphi_i^*=0)$ is the F-statistic associated with short-run causality from real exchange rates to real oil prices

To summarize, the post-1970 data provides a rather weak support in favor of Dutch disease hypothesis. Evidence of significant long-run effects from oil prices and exchange rates is limited to only three countries of Bolivia, Mexico and Norway. Incidentally, these countries have operated under a crawling peg (Bolivia) or an independent float (Mexico and Norway) exchange rate regime in the recent past. For these countries, we find that high oil prices contributed to real exchange rate appreciation in the long-run. In contrast, the remaining ten countries have operated under a managed float (Algeria, Colombia, Gabon, Indonesia and Nigeria) or a fixed-peg (Kuwait, Russia, Saudi Arabia, Venezuela and Angola) regime. Thus, countries with more flexible exchange rate regimes appear more susceptible to Dutch disease phenomena.

Nevertheless, these findings provide a weaker assessment of the effect of oil prices on exchange rates and the possibility of Dutch disease than those reported in Issa et al. (2008), Jahan-Parvar and Mohammadi (2009, 2010), Koranchelian (2005), Zalduendo (2006), and Kalcheva and Oomes (2007).

The weak oil price—real exchange rate relationship evident in most countries implies that most oil exporting countries have been able to successfully isolate their domestic economy by allocating the resulting foreign exchange earnings directly to imports, or through investments abroad via the creation of sovereign wealth funds. As for the three remaining countries, the strong oil price—exchange rate relationship has a number of policy implications. First, if high oil prices and gains in terms of trade are expected to be temporary then appropriate policy requires the protection of vulnerable industries. However, if the gains in terms of trade are permanent, then appropriate policy requires major structural adjustments which facilitate the reallocation of resources from the traditional exports sector to non-tradables, as well as the diversification of the exports sector to make them less vulnerable to external shocks.

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