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# Oil prices and real exchange rates

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**Abstract**

In this paper, we investigate the long-run relationship between real oil prices and real exchange rates by using a monthly panel of G7 countries from 1972:1 to 2005:10. We first test whether exchange rates are cointegrated with real oil prices. It is shown that real oil prices may have been the dominant source of real exchange rate movements and that there is a link between real oil prices and real exchange rates. We then examine the ability of real oil prices to forecast future real exchange returns. Panel predictive regression estimates suggest that real oil prices have significant forecasting power. The out-of-sample prediction performances demonstrate greater predictability over longer horizons.


*JEL classification:* Q43; F31; F47

*Keywords:* Real oil prices; Real exchange rates; Prediction; Long horizons; Panel cointegration

## 1. Introduction

It is well known that real exchange rate fluctuations can be attributed primarily to non-monetary shocks. Clarida and Gali (1994) use the Blanchard–Quah identification strategy to estimate the share of exchange rate variability that is due to different shocks by using quarterly US–Canada, US–Germany, US–Japan, and US–UK real exchange rate data from 1974:Q3 to 1992:Q4. They find that real shocks can account for more than 50% of the variance of real exchange rate changes over all time horizons. Lastrapes (1992) also uses the Blanchard–Quah

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approach to estimate structural VARs and obtains similar results to those of Clarida and Gali (1994).[2]

Different sources of real shocks have been investigated in Zhou (1995). Among many sources of real disturbances, such as oil prices, fiscal policy, and productivity shocks, it has been shown that oil price fluctuations play a major role in explaining real exchange rate movements. Moreover, Chaudhuri and Daniel (1998) investigate 16 OECD countries and find that the nonstationary behavior of US dollar real exchange rates is due to the nonstationary behavior of real oil prices. Similar results are obtained by Amano and Norden (1998a,b). By using data on real effective exchange rates for Germany, Japan, and the US, they find that the real oil price is the most important factor determining real exchange rates in the long run. Camarero and Tamarit (2002) use panel cointegration techniques to investigate the relationship between real oil prices and the Spanish peseta's real exchange rate.

In this paper, we study the nexus between real oil prices and real exchange rates for a sample of G7 countries by using monthly panel data from 1972:1 to 2005:10. By using a panel study, we first reexamine whether exchange rates are cointegrated with real oil prices, as has been documented in the existing literature, which comprises mainly time-series case studies of individual countries. That is, we investigate whether pooling across currencies and using panel cointegration techniques generate different implications for real oil prices and real exchange rates. The second issue is to examine the ability of real oil prices to forecast future exchange rate returns in line with research from prediction regression tests that investigate the relationship between exchange rates and fundamentals. See, for example, Mark (1995), Kilian (1999), Mark and Choi (1997), and Mark and Sul (2001).

This paper contributes to the literature in three distinctive ways. First, we consider different measures of oil prices, including the world price of oil, the United Arab Emirates price of oil (Dubai), the British price of oil (Brent), and the US West Texas Intermediate price of oil (WTI); in previous work, only one of these oil prices has been used. Second, we pool the data and apply tests for panel unit roots and heterogeneous panel cointegration, which may help improve the power of the tests. Most importantly, to the best of our knowledge, this is the first paper to assess the role of real oil prices in predicting real exchange rates over long horizons. Although other real factors, such as productivity differentials and real interest rate differentials, have been investigated previously (see Mark and Choi, 1997), no study has yet focused on the long-horizon forecasting power of real oil prices. In this paper, we fill this gap and explore the ability of real oil prices to explain movements in real exchange rates in a predictive regression framework.

The rest of the paper is organized as follows. In Section 2, we present a simple theoretical model. In Section 3, we describe the sources of data and report our country-by-country test results for unit roots and cointegration. In Section 4, we report the panel test results. In Section 5, we use the prediction regression test to determine whether the real oil price is able to predict real exchange rate movements. Out-of-sample forecasting performance is also examined. In Section 5, we also present robustness checks. Concluding remarks are provided in Section 6.

## 2. A simple theoretical model

In this section, we present a simple theoretical model to motivate our investigation of the link between real exchange rates and oil prices. Suppose that both traded and nontraded goods are

[2] Clarida and Gali (1994) identify three shocks: supply shocks, demand shocks, and monetary shocks, whereas Lastrapes (1992) only identifies two shocks, a real shock and a nominal shock.

produced in the home and foreign countries. Let the log-linear approximation of the home and foreign country consumer price indexes be

$$p = \alpha p^{\mathrm{T}} + (1-\alpha)p^{\mathrm{N}}, \tag{1}$$

$$p = \alpha^* p^{\mathrm{T}*} + (1-\alpha^*)p^{\mathrm{N}*}, \tag{2}$$

where $p^{\mathrm{T}}$ ($p^{\mathrm{T}*}$) and $p^{\mathrm{N}}$ ($p^{\mathrm{N}*}$) are prices of traded and nontraded goods in the home (foreign) country, respectively. The $\alpha$ and $\alpha^*$ weights correspond to the expenditure shares on traded goods near the point of approximation for the home and foreign countries, respectively. The log of the real exchange rate, $q_t$, is defined as

$$q_t = s + p^* - p, \tag{3}$$

where $s$ is the log of the nominal exchange rate (the domestic currency price of foreign exchange). Thus, the real exchange rate can be rewritten as

$$q_t = (s + p^{\mathrm{T}*} - p^{\mathrm{T}}) + (1-\alpha)(p^{\mathrm{T}} - p^{\mathrm{N}}) - (1-\alpha^*)(p^{\mathrm{T}*} - p^{\mathrm{N}*}). \tag{4}$$

According to Eq. (4), if $\alpha \simeq \alpha^*$, a rise in the relative price of domestic tradables, $(p^{\mathrm{T}} - p^{\mathrm{N}})$, depreciates the real exchange rate, while the magnitude of the rise exceeds that of the rise in the relative price of foreign tradables, $(p^{\mathrm{T}*} - p^{\mathrm{N}*})$. That is, if the home country is more dependent on imported oil, a real oil price rise may increase the prices of tradable goods in the home country by a greater proportion than in the foreign country, and thereby cause a real depreciation of the home currency. Moreover, in order to improve competitiveness when an oil price shock worsens the term of trade, the home country would have to raise the nominal exchange rate, which would lead to a further real depreciation.

In this paper, we empirically investigate the link between real exchange rates and real oil prices. We study whether the real exchange rate is positively related to the real oil price as predicted by the above model. We also investigate whether real oil prices can predict movements in real exchange rates.[3]

## 3. The data and a reexamination of country-by-country results

We use data on the G7 countries, Canada, France, Germany, Italy, Japan, the UK, and the US. Monthly data, mainly from 1972:1 to 2005:10, are used. Data on nominal exchange rates (domestic currency per unit of foreign currency), consumer price indexes, the world price of oil, the United Arab Emirates price of oil, the British price of oil, and the WTI were obtained from International Financial Statistics (IFS), published by the International Monetary Fund (IMF). Real

[3] As pointed out by an anonymous referee, the empirical results may depend on the dependence on imported oil. This dependence is important in explaining movements in real exchange rates in response to oil price shocks. Some studies deal with heterogeneity in the degree of dependence on oil. See Camarero and Tamarit (2002). However, as is explained in the context of the model, attempts to improve balance of trade positions also explain movements in real exchange rates. In this paper, because we intend to investigate the overall effects of oil prices on real exchange rates, we do not try to disentangle these effects.

Table 1
Data

| Data | Source | Code |
|---|---|---|
| Nominal exchange rate | IFS | 156..AE.ZF..., 132..AE.ZF..., 134..AE.ZF..., 136..AE.ZF..., 158..AE.ZF..., 112..AG.ZF... |
| Consumer price index | IFS | 15664...ZF..., 13264...ZF..., 13464.D.F (West Germany, 1972:1–1991:12), 13464...ZF... (Unified Germany, 1991:1–2005:10), 13664...ZF..., 15864... ZF..., 11264...ZF..., 11163BA.ZF... |
| World average crude price | IFS | 00176AAZZF... |
| Dubai | IFS | 46676AAZZF... |
| UK Brent | IFS | 11276AAZZF... |
| West Texas Intermediate | IFS | 11176AAZZFM17 |
| Real GDP | DATASTREAM | USOCFGDPD, JPOCFGDPD, UKOCFGDPD, BDOCFGDPD, ITOCFGDPD, CNOCFGDPD, FROCFGDPD |
| Employment | DATASTREAM | BDOEM020O, FROEM020O, ITI67E..O, USQ67E..O |
| | IFS | 11267..CZF..., 15667EY.ZF..., 15867EYCZF... |
| Nominal interest rate | IFS | 15660C..ZF..., 13260C..ZF..., 13460B..ZF..., 13660B..ZF..., 15860B..ZF..., 11260C..ZF..., 11160C..ZF... |

exchange rates are constructed by using consumer price indexes, and real oil prices are defined as the US dollar prices of oil converted to the domestic currency and then deflated by the domestic consumer price index (CPI). CPI data are available from 1972:1 to 1991:12 for West Germany and from 1991:1 to 2005:10 for Germany. We therefore combined these two series. Data on the WTI are only available from 1983:1 to 2005:10. All variables are measured in logarithms and the US is chosen to be the numeraire country. Variable names and data codes are provided in Table 1.

Before we investigate the stationarity of these time series, we first implement tests of structural breaks. We do so because oil prices and exchange rates may have been affected by major events. Failing to account for structural changes may bias the tests for stationarity. We apply the tests proposed by Bai and Perron (1998, 2003), which allow for multiple structural changes.[4] As suggested by Bai and Perron (2003), a useful strategy is to first examine the double maximum $UD_{max}$ or $WD_{max}$ tests to see if at least one break is present given some upper bound, $M$. The null hypothesis is that there is no structural break, and the alternative is that there are an unknown number of breaks. The maximum number of structural changes allowed was chosen to be five ($M=5$). Test results are reported in a Supplementary material.[5] According to the structural break test results from the $UD_{max}$ or $WD_{max}$ tests, for most of the series, the null hypothesis of no structural break is not rejected at the 5% significant level. As exceptions, the $WD_{max}$ statistics for real exchange rates in France and Germany reject the null of no structural break at the 5% level but not at the 1% significant level.[6]

Given no evidence of the presence of structural breaks, we test for unit roots. We examine each individual series by using the Augmented Dickey–Fuller (ADF) test as well as the four modified

[4] The GAUSS code is available at: http://qed.econ.queensu.ca/jae/2003-v18.1/bai-perron/.

[5] To save space, all the results in this section (for tests of structural breaks, tests of unit roots, and tests of cointegration) are reported in a Supplementary material, see Appendix A.

[6] The $WD_{max}$ statistics are 10.4671 and 10.9996 for French and German real exchange rates, respectively. The critical values are 9.9100 (5%) and 13.8300 (1%).

Table 2
Panel unit root tests

| | Real exchange rates | Real oil prices | | | |
|---|---|---|---|---|---|
| | | World prices | Dubai | Brent | WTI |
| *Series in level* | | | | | |
| Levin, Lin, and Chu | 0.47 (0.68) | −3.26 (0.00) | −2.54 (0.01) | −1.30 (0.10) | −0.04 (0.49) |
| Breitung | 1.49 (0.93) | −0.63 (0.26) | −0.63 (0.26) | −0.19 (0.42) | 1.16 (0.88) |
| Im, Pesaran, and Shin | 0.42 (0.66) | −1.91 (0.03) | −1.71 (0.04) | −1.07 (0.14) | 0.76 (0.78) |
| Fisher–ADF | 8.21 (0.77) | 19.19 (0.08) | 18.20 (0.11) | 14.30 (0.28) | 5.73 (0.93) |
| Fisher–PP | 8.88 (0.71) | 15.34 (0.22) | 14.48 (0.27) | 12.87 (0.38) | 3.98 (0.98) |
| *Series in first differences* | | | | | |
| Levin, Lin, and Chu | −73.01 (0.00) | −65.91 (0.00) | −78.69 (0.00) | −79.04 (0.00) | −40.25 (0.00) |
| Breitung | −32.77 (0.00) | −29.57 (0.00) | −39.38 (0.00) | −38.60 (0.00) | −21.31 (0.00) |
| Im, Pesaran, and Shin | −51.05 (0.00) | −39.4259 (0.00) | −49.2408 (0.00) | −50.4664 (0.00) | −30.1655 (0.00) |
| Fisher–ADF | 1059.23 (0.00) | 813.969 (0.00) | 1037.1 (0.00) | 1057.84 (0.00) | 551.99 (0.00) |
| Fisher–PP | 1101.45 (0.00) | 1002.15 (0.00) | 1030.07 (0.00) | 1054.21 (0.00) | 632.672 (0.00) |

Note: The null hypothesis is that the series is a unit root process. An intercept and trend are included in the test equation. *p*-values are provided in parentheses. Probabilities for Fisher-type tests were computed by using an asymptotic $\chi^2$ distribution. All other tests assume asymptotic normality. The lag length was selected by using the Akaike Information Criteria.

tests (*M*-tests) proposed by Ng and Perron (2001) based on modified information criteria (MIC): the modified Phillips–Perron test ($MZ_{\alpha}$); the modified Sargan–Bhargava test (MSB); the modified point optimal test ($MP_T$); and the modified Phillips–Perron $MZ_t$ test ($MZ_t$).[7] Clearly, according to the ADF tests, all real exchange rates are integrated of order one ($I(1)$). This result is consistent with well-documented evidence of the nonstationary behavior of real exchange rates from the literature. It is also evident that real oil prices are nonstationary. This confirms the findings of Amano and Norden (1998a), who have shown that German and Japanese real WTI prices are $I(1)$ series. These results are also consistent with those of Chaudhuri and Daniel (1998), who find that real Dubai oil prices for Canada, Germany, Italy, Japan, and the UK are $I(1)$. It is worth noting that results are mixed for real world oil prices, Brent oil prices, and the WTI when Ng and Perron's (2001) M-tests are used; however, overall, the evidence suggests the nonstationarity of real exchange rates and real Dubai oil prices.

We apply the Johansen test for cointegration and investigate the results from the trace and max-eigenvalue statistics.[8] In most cases, real exchange rates and real oil prices are not cointegrated in Canada, France, Italy, and the UK. On the other hand, for the German mark and the Japanese yen, both the trace and the max-eigenvalue tests suggest a cointegrating relationship between oil prices and real exchange rates. These mixed results on cointegration are consistent with the findings of Amano and Norden (1998a) and Chaudhuri and Daniel (1998).[9] The evidence may be mixed because of a lack of power in the individual country-by-country tests.

[7] The GAUSS code is available at: http://www-personal.umich.edu/~ngse/research.html.

[8] A lag of four was chosen, but the test results are similar when applying different lag structures. We assume that the levels data have no deterministic trends and that the cointegrating equations have intercepts.

[9] Amano and Norden (1998a) use the Johansen test whereas Chaudhuri and Daniel (1998) apply the Engle–Granger test.

Table 3
Pedroni (2004) panel cointegration tests

| | World prices | | Dubai | | Brent | | WTI | |
|---|---|---|---|---|---|---|---|---|
| | No trend | Trend | No trend | Trend | No trend | Trend | No trend | Trend |
| Panel $v$ | 1.72* | 1.19 | 2.38* | 1.39* | 2.00* | 1.41* | 2.11* | 0.35 |
| Panel $\rho$ | −2.16* | −1.13 | −2.82* | −1.42* | −2.53* | −1.47* | −1.21 | −0.41 |
| Panel PP | −2.34* | −1.22 | −2.39* | −1.30* | −2.35* | −1.37* | −0.39 | −0.10 |
| Panel ADF | −2.12* | −0.94 | −2.35* | −1.16 | −2.35* | −1.30* | −1.05 | −0.59 |
| Group $\rho$ | −0.79 | −0.09 | −1.40* | −0.36 | −1.16 | −0.37 | −0.32 | −0.29 |
| Group PP | −1.79* | −0.70 | −1.88* | −0.81 | −1.85* | −0.85 | 0.21 | −0.06 |
| Group ADF | −1.58* | −0.45 | −1.84* | −0.72 | −1.82* | −0.85 | −0.70 | −0.77 |

Note: The null hypothesis is that there is no cointegration. An asterisk (*) indicates rejection at the 10% level or better.

To sum up, a preliminary investigation of the data yields evidence that is consistent with existing country-by-country results. Our findings are robust to using different measures of real oil prices.

## 4. Panel results

### *4.1. Panel tests of unit roots and cointegration*

As shown in the previous section, our country-by-country results are consistent with other empirical studies. Given the notoriously low power of individual country-by-country tests for unit roots and cointegration, it may be preferable to pool the currencies and conduct panel analysis. We implement three different types of panel unit root tests: the Levin et al. (2002) test (LLC); the Breitung (2000) test; the Im et al. (2003) test (IPS); and the Fisher-type ADF and Phillips–Perron (PP) tests (see Maddala and Wu (1999)). Table 2 reports the test results based on the inclusion of an intercept and trend. It is clear that the real exchange rate is an $I(1)$ series. The evidence for oil prices is relatively weak because the LLC and IPS tests suggest stationarity in some cases. However, significant evidence of nonstationarity is suggested by the Fisher-type statistics (Fisher–PP), which have been shown to be superior to the LLC and IPS tests by Maddala and Wu (1999).

We then implemented the panel cointegration tests proposed by Pedroni (2004). This is a residual-based test for the null of no cointegration in heterogeneous panels. Two classes of statistics are considered in the context of the Pedroni test. The first type is based on pooling the residuals of the regression along the within-dimension of the panel, whereas the second type is based on pooling the residuals of the regression along the between-dimension of the panel. For the first type, the test statistics are the panel $v$-statistic, the panel $\rho$-statistic, the panel PP-statistic, and the panel ADF-statistic. These statistics are constructed by taking the ratio of the sum of the numerators and the sum of the denominators of the analogous conventional time-series statistics across the individual members of the panel. The tests for the second type include the group $\rho$-statistic, the group PP-statistic, and the group ADF-statistic. They are simply the group mean statistics of the conventional individual time-series statistics. All statistics have been standardized by the means and variances so that they are asymptotically distributed $N(0,1)$ under the null of no cointegration. As one-sided tests, large positive values of the panel $v$-statistic reject the null hypothesis of no cointegration. For the remaining statistics (the panel $\rho$, the panel PP, the panel ADF, the group $\rho$, the group PP, and the group ADF tests), large negative values reject this null. See Pedroni (2004) for a detailed discussion.

Table 4
Larsson et al. (2001) panel cointegration tests

| | Standardized LR-bar | |
|---|---|---|
| | $r=0$ | $r=1$ |
| World prices | 3.999 | 1.695 |
| Dubai | 3.789 | 1.677 |
| Brent | 3.403 | 1.732 |
| WTI | 1.994 | 1.980 |

Note: The null hypothesis is that there are no more than $r$ cointegrating relationships. Critical values are 1.29 (10%), 1.64 (5%) and 2.32 (1%).

Table 3 reports the results of the Pedroni cointegration tests. Clearly, the panel statistics indicate fairly strong support for the hypothesis that oil prices are cointegrated with real exchange rates in the context of world prices, Dubai prices, and Brent prices whereas the evidence is weaker for WTI prices. Nevertheless, evidence obtained from Pedroni's (2004) test seems to suggest a long-run equilibrium relationship between real oil prices and real exchange rates.

For consistency with the country-by-country Johansen cointegration tests of the previous section, we implemented the likelihood-based cointegration test proposed by Larsson et al. (2001). To allow for the possibility of multiple cointegrating vectors, Larsson et al. (2001) propose a likelihood-based test of the cointegrating rank in heterogeneous panels. Under the null hypothesis, each group in the panel has at most $r$ cointegrating relationships. After obtaining the average of the individual Johansen trace statistics (namely the LR-bar statistic), we derived a standardized LR-bar statistic to use as the basis for the panel cointegration rank test. The asymptotic distribution of the standardized LR-bar is standard normal. See Larsson et al. (2001) for a detailed discussion. We report the results from the Larsson et al. (2001) test in Table 4. Because the test follows a standard normal distribution, its 5% and 1% critical values are 1.645 and 2.325, respectively. Clearly, the results suggest either one (or two) cointegrating vectors between real oil prices and real exchange rates at the 1% level (or the 5% level). Compared with the Pedroni tests, Larsson et al.'s (2001) test provides stronger evidence of cointegration.

Table 5
Pesaran (2006) CIPS test statistics

| Lag | $p=1$ | $p=2$ | $p=3$ | $p=4$ |
|---|---|---|---|---|
| *Series in levels* | | | | |
| Real exchange rates | −1.43 | −1.50 | −1.64 | −1.68 |
| World prices | −2.16 | −2.38 | −2.54 | −2.58 |
| Dubai | −2.18 | −2.39 | −2.55 | −2.57 |
| Brent | −2.16 | −2.44 | −2.54 | −2.58 |
| WTI | −2.60 | −2.64 | −2.83 | −2.73 |
| *Series in first differences* | | | | |
| Real exchange rates | −13.44 | −10.63 | −9.32 | −8.32 |
| World prices | −12.98 | −10.72 | −9.45 | −8.24 |
| Dubai | −13.02 | −10.67 | −9.45 | −8.20 |
| Brent | −12.97 | −10.66 | −9.43 | −8.17 |
| WTI | −10.81 | −9.05 | −8.01 | −7.05 |

Critical values for the CIPS test are −2.73 (10%), −2.83 (5%), and −3.03 (1%). See Table 2c in Pesaran (2006).

Table 6
Mark and Sul (2001) panel dynamic OLS-based cointegration tests

| | $\theta_{\text{dols}}$ | $t$-ratio | $p$-value[a] | $p$-value[b] |
|---|---|---|---|---|
| World oil prices | −0.0263 | −5.7510 | 0.0100 | 0.0310 |
| Dubai oil prices | −0.0260 | −5.7890 | 0.0140 | 0.0300 |
| Brent oil prices | −0.0273 | −5.8150 | 0.0100 | 0.0270 |
| WTI oil prices | −0.0401 | −5.5260 | 0.0330 | 0.0690 |

[a] $p$-values from parametric bootstrap.
[b] $p$-values from nonparametric bootstrap.

In summary, in contrast to the mixed evidence from the country-by-country tests, the panel results from the residual-based and likelihood-based tests show that oil prices can adequately capture permanent innovations in the real exchange rate in the long run.

### 4.2. Cross-section dependence

The panel unit root tests and cointegration tests applied in the previous section do not account for cross-sectional dependence of the contemporaneous error terms. It has been shown in the literature that failing to consider cross-sectional dependence may cause substantial size distortions in panel unit root tests. See, for example, O'Connell (1998) and Pesaran (2006). To check the robustness of the order of integration and the existence of cointegration, in this section, we implement panel tests of unit roots and cointegration that account for the presence of cross-section dependence.

We apply the *CIPS* test of panel unit roots proposed by Pesaran (2006). We first obtain the $t$-ratio of the OLS estimate of $b_i$ in the following cross-sectionally augmented DF (CADF) regression:

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + c_i \overline{y}_{t-1} + \sum_{j=0}^{p} d_{ij} \Delta \overline{y}_{t-j} + \sum_{j=1}^{p} k_{ij} \Delta y_{i,t-j} + e_{ij}.$$

We denote the individual $t$-ratio by $\text{CADF}_i$. To test the null hypothesis of a unit root, we compute the cross-sectionally augmented version of the IPS test as $\text{CIPS} = N^{-1} \sum_{i=1}^{N} \text{CADF}_i$. An intercept and trend are included. Table 5 reports the Pesaran (2006) CIPS statistics with $p=1, 2, 3,$ and 4.[10] It is clear that, after accounting for cross-section dependence, at the 5% significant level, the finding that real exchange rates and real oil prices are $I(1)$ is confirmed.

Next, we implement the panel unit root test proposed by Mark and Sul (2001) in the following regression:

$$\Delta y_{it} = \gamma_{it} + \theta w_{it} + \varepsilon_{it},$$

where $w_{it} = y_{it-1} - \beta x_{it-1}$. Under the null hypothesis that $w_{it}$ is nonstationary, a test of $\theta = 0$ is a test of the hypothesis that $y_t$ and $x_t$ are not cointegrated. To account for cross-sectional dependence,

[10] There were no substantial changes in results for greater values of $p$.

Table 7
Panel estimates of the cointegrating relationship: FMOLS/DOLS in Pedroni (2000, 2001) and PMG in Pesaran et al. (1999)

| Country | World oil prices | | | | | | Dubai oil prices | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | FMOLS | | DOLS | | PMG | | FMOLS | | DOLS | | PMG | |
| | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics |
| Canada | 0.11 | 6.32 | 0.12 | 5.80 | 0.12 | 1.34 | 0.12 | 7.08 | 0.12 | 6.66 | 0.12 | 1.54 |
| France | 0.20 | 5.84 | 0.21 | 3.60 | 0.35 | 2.84 | 0.20 | 6.19 | 0.21 | 3.93 | 0.30 | 2.73 |
| Germany | 0.22 | 6.59 | 0.24 | 4.07 | 0.40 | 3.44 | 0.22 | 6.94 | 0.23 | 4.47 | 0.34 | 3.39 |
| Italy | 0.24 | 7.07 | 0.25 | 4.97 | 0.34 | 2.48 | 0.24 | 7.72 | 0.25 | 5.58 | 0.29 | 2.40 |
| Japan | 0.38 | 8.71 | 0.40 | 8.04 | 0.47 | 4.18 | 0.38 | 9.13 | 0.39 | 8.90 | 0.43 | 3.86 |
| UK | 0.21 | 4.72 | 0.50 | 9.81 | 0.28 | 1.30 | 0.21 | 5.06 | 0.48 | 10.00 | 0.23 | 1.12 |
| **Panel** | 0.23 | 16.02 | 0.29 | 14.81 | 0.38 | 6.44 | 0.23 | 17.19 | 0.28 | 16.14 | 0.33 | 6.04 |

Note: The regression is $q_{it}=\alpha_i+\beta_i p_{it}+\varepsilon_{it}$, where $q_{it}$ is the real exchange rate and $p_{it}$ is the real oil price.

Mark and Sul (2001) propose using both parametric and nonparametric bootstrap $p$-values to test the null of no cointegration. In the spirit of Maddala and Wu (1999), the bootstrapping procedure preserves the cross-sectional dependence exhibited by the estimated residuals. The coefficients $\beta$ and $\theta$ are estimated by using a panel dynamic OLS estimator. Clearly, both the asymptotic test and the bootstrapping $p$-values reported in Table 6 comfortably reject the null hypothesis of no cointegration between the real exchange rate and the real oil price for all measures of oil prices.

To sum up, after accounting for the cross-sectional dependence of the contemporaneous error terms, the findings relating orders of integration and cointegration between real exchange rates and real oil prices are confirmed; hence, our results are robust.

## 5. The long-run equilibrium exchange rate and the oil price

Given the evidence that real exchange rates and real oil prices may be cointegrated, we thus estimate the cointegrating coefficients to investigate the long-run relationship between them. We apply three estimation methods: between-dimension panel fully modified OLS (FMOLS); between-dimension panel dynamic OLS (DOLS); and pooled mean group estimation (PMG).

Table 8
Panel estimates of the cointegrating relationship: FMOLS/DOLS in Pedroni (2000, 2001) and PMG in Pesaran et al. (1999)

| Country | World oil prices | | | | | | Dubai oil prices | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | FMOLS | | DOLS | | PMG | | FMOLS | | DOLS | | PMG | |
| | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics | $\beta$ | t-statistics |
| Canada | 0.12 | 6.59 | 0.13 | 6.52 | 0.12 | 1.41 | 0.11 | 2.95 | 0.13 | 2.50 | 0.09 | 0.42 |
| France | 0.23 | 6.86 | 0.24 | 4.71 | 0.33 | 2.99 | 0.37 | 9.41 | 0.39 | 7.27 | 0.38 | 2.92 |
| Germany | 0.25 | 7.63 | 0.26 | 5.35 | 0.37 | 3.66 | 0.39 | 9.88 | 0.40 | 7.67 | 0.42 | 3.40 |
| Italy | 0.27 | 8.44 | 0.28 | 6.85 | 0.32 | 2.66 | 0.36 | 8.25 | 0.37 | 6.43 | 0.38 | 2.46 |
| Japan | 0.42 | 10.94 | 0.43 | 12.05 | 0.46 | 4.90 | 0.44 | 12.82 | 0.45 | 9.48 | 0.44 | 6.13 |
| UK | 0.25 | 6.01 | 0.52 | 10.56 | 0.29 | 1.45 | 0.34 | 6.99 | 0.23 | 6.39 | 0.42 | 2.31 |
| **Panel** | 0.25 | 18.97 | 0.31 | 18.79 | 0.37 | 6.93 | 0.34 | 20.53 | 0.33 | 16.23 | 0.41 | 8.22 |

Note: The regression is $q_{it}=\alpha_i+\beta_i p_{it}+\varepsilon_{it}$, where $q_{it}$ is the real exchange rate and $p_{it}$ is the real oil price.

Table 9
LSDV estimates and bootstrapped $t$-ratios for panel predictive regression

| | $\delta_{\mathrm{LSDV}}$ | $t$-ratio | $p$-value[a] | $p$-value[b] |
|---|---|---|---|---|
| (1) World prices | −0.0222 | −5.1760 | 0.0010 | 0.0000 |
| (2) Dubai | −0.0223 | −5.2140 | 0.0010 | 0.0000 |
| (3) Brent | −0.0229 | −5.2120 | 0.0010 | 0.0010 |
| (4) WTI | −0.0375 | −5.0570 | 0.0020 | 0.0010 |

Note: The model is $q_{it+k}-q_{it}=\gamma_{ik}+\delta_{ik}z_{it}+\varepsilon_{it+k}$, where $q_{it}$ is the log of the real exchange rate and $z_{it}$ is the deviation of the log of the real exchange rate from its long-run equilibrium value.

[a] $p$-values from parametric bootstrap.

[b] $p$-values from nonparametric bootstrap.

FMOLS and DOLS were proposed by Pedroni (2000, 2001), and PMG is suggested by Pesaran et al. (1999). Consider the following regression:

$$q_{it} = \alpha_i + \beta p_{it} + \varepsilon_{it}, \tag{5}$$

where $q_{it}$ is the log of the real exchange rate and $p_{it}$ is the log of the real oil price. Tables 7 and 8 present the estimation results for Eq. (5) based on FMOLS, DOLS, and PMG. Individual estimates and $t$-statistics for $H_0$: $\beta_i=0$ are reported in the first six entries. At the bottom of the table, panel estimation results are reported.

In Tables 7 and 8, most of the slope coefficients are statistically significant. The panel estimates are highly significant for all measures of the oil price and for all estimators (FMOLS, DOLS, and PMG). According to the panel estimates, a rise in real oil prices depreciates the real exchange rates in the long run. This result is consistent with the arguments regarding the link between the price of oil and exchange rate dynamics based on the theoretical model. First, a real oil price rise may increase the prices of tradables relative to nontradables in both the domestic country and in the US. If a domestic country is more dependent on imported oil than is the US, the increase in the price of tradables relative to nontradables in the domestic country would exceed the US increase and thus cause a real depreciation of the domestic currency against the US dollar. Furthermore, to improve competitiveness when an oil shock worsens the term of trade, the home country would have to raise the nominal exchange rate, which would cause a further real depreciation.

## 6. Panel prediction regression

In this section, we investigate the relationship between real exchange rates and real oil prices by using the prediction regression test proposed by Mark (1995) and Mark and Choi (1997). We focus on whether deviations of real exchange rates from real oil prices (represented by the long-run equilibrium error) incorporate predictive content for future movements of real exchange rates. Consider prediction regressions of the form:

$$q_{it+k}-q_{it} = \gamma_{ik} + \delta_k z_{it} + \varepsilon_{it+k}, \tag{6}$$

where $z_{it}$ is the deviation of the real exchange rate from its long-run value predicted by the cointegrating relationship in Eq. (5). That is,

$$z_{it} = q_{it}-\hat{\beta} p_{it}, \tag{7}$$

Table 10
Out-of-sample forecasts of real exchange rate returns

| Country | World prices | | | Dubai | | |
|---|---|---|---|---|---|---|
| | $U$ | p-value [a] | $p$-value [b] | $U$ | p-value [a] | $p$-value [b] |
| *A. One-month-ahead forecasts* | | | | | | |
| Canada | 1.006 | 0.831 | 0.746 | 1.006 | 0.810 | 0.737 |
| France | 0.991 | 0.000 | 0.002 | 0.991 | 0.002 | 0.003 |
| Germany | 0.991 | 0.001 | 0.002 | 0.992 | 0.006 | 0.004 |
| Italy | 0.988 | 0.000 | 0.000 | 0.988 | 0.001 | 0.001 |
| Japan | 0.991 | 0.006 | 0.003 | 0.991 | 0.006 | 0.006 |
| UK | 0.993 | 0.014 | 0.018 | 0.993 | 0.019 | 0.016 |
| *B. Sixteen-month-ahead forecasts* | | | | | | |
| Canada | 1.021 | 0.555 | 0.521 | 1.047 | 0.662 | 0.620 |
| France | 0.947 | 0.062 | 0.074 | 0.936 | 0.046 | 0.055 |
| Germany | 0.960 | 0.089 | 0.105 | 0.946 | 0.064 | 0.069 |
| Italy | 0.887 | 0.006 | 0.010 | 0.884 | 0.006 | 0.010 |
| Japan | 1.016 | 0.606 | 0.544 | 0.992 | 0.329 | 0.315 |
| UK | 1.048 | 0.783 | 0.699 | 1.033 | 0.690 | 0.611 |
| *C. Twenty-four-month-ahead forecasts* | | | | | | |
| Canada | 0.989 | 0.325 | 0.322 | 1.019 | 0.481 | 0.445 |
| France | 0.921 | 0.059 | 0.080 | 0.910 | 0.054 | 0.057 |
| Germany | 0.940 | 0.100 | 0.113 | 0.924 | 0.071 | 0.078 |
| Italy | 0.825 | 0.006 | 0.007 | 0.825 | 0.007 | 0.010 |
| Japan | 0.998 | 0.421 | 0.386 | 0.967 | 0.226 | 0.241 |
| UK | 1.039 | 0.673 | 0.572 | 1.027 | 0.594 | 0.509 |

Note: $U$ is Theil's $U$-statistic.
[a] $p$-values from parametric bootstrap.
[b] $p$-values from nonparametric bootstrap.

where $\hat{\beta}$ is the panel FMOLS estimate of $\beta$ reported in Tables 7 and 8.[11] If $q_{it}$ is greater than its long-run equilibrium value, $\beta p_{it}$, the currency is predicted to appreciate. Furthermore, if the real oil price helps forecast the real exchange rate, $\delta_{ik}<0$ should be significantly different from zero.

## 6.1. Panel estimation of short-horizon prediction

Following Mark and Sul (2001), we initially focus on short-horizon prediction ($k=1$). To estimate Eq. (6), we simply apply least-squares-with-dummy-variables (LSDV) estimators and bootstrap the $t$-ratio to correct for the small sample bias. Both parametric and nonparametric bootstraps are employed as proposed by Mark and Sul (2001).[12] Table 9 reports the LSDV estimates, the $t$-statistics, and the $p$-values for the bootstrapped $t$-statistics. Clearly, the sign of the coefficient is negative and is consistent with the prediction of long-run equilibrium error correction. Moreover, all tests significantly reject the hypothesis that real oil prices have no

[11] Results based on panel DOLS and PMG were similar.
[12] We apply the bootstrap with 1000 replications. See Mark and Sul (2001) for details of the bootstrapping procedure and the data-generating process used in the bootstrap. The bootstrapping procedure preserves the cross-sectional dependence exhibited in the estimated residuals. The GAUSS code is available on the web page of Nelson Mark: http://www.nd.edu/~nmark/.

Table 11
Out-of-sample forecasts of real exchange rate returns

| Country | Brent | | | WTI | | |
|---|---|---|---|---|---|---|
| | $U$ | p-value [a] | $p$-value [b] | $U$ | p-value[a] | $p$-value[b] |
| *A. One-month-ahead forecasts* | | | | | | |
| Canada | 1.010 | 0.870 | 0.778 | 1.042 | 0.820 | 0.788 |
| France | 0.991 | 0.002 | 0.003 | 0.992 | 0.041 | 0.048 |
| Germany | 0.991 | 0.004 | 0.004 | 0.988 | 0.012 | 0.014 |
| Italy | 0.989 | 0.001 | 0.000 | 1.000 | 0.393 | 0.362 |
| Japan | 0.989 | 0.004 | 0.002 | 0.970 | 0.000 | 0.000 |
| UK | 0.993 | 0.016 | 0.014 | 1.008 | 0.722 | 0.702 |
| *B. Sixteen-month-ahead forecasts* | | | | | | |
| Canada | 1.110 | 0.791 | 0.727 | 1.689 | 0.852 | 0.856 |
| France | 0.952 | 0.086 | 0.083 | 1.065 | 0.643 | 0.620 |
| Germany | 0.959 | 0.097 | 0.104 | 1.033 | 0.543 | 0.507 |
| Italy | 0.895 | 0.011 | 0.011 | 1.148 | 0.799 | 0.783 |
| Japan | 0.974 | 0.205 | 0.207 | 0.597 | 0.000 | 0.000 |
| UK | 1.019 | 0.605 | 0.487 | 1.336 | 0.906 | 0.906 |
| *C. Twenty-four-month-ahead forecasts* | | | | | | |
| Canada | 1.098 | 0.694 | 0.624 | 1.751 | 0.799 | 0.797 |
| France | 0.936 | 0.104 | 0.107 | 1.097 | 0.641 | 0.621 |
| Germany | 0.948 | 0.142 | 0.126 | 1.065 | 0.583 | 0.562 |
| Italy | 0.843 | 0.008 | 0.013 | 1.181 | 0.754 | 0.744 |
| Japan | 0.960 | 0.197 | 0.207 | 0.697 | 0.004 | 0.005 |
| UK | 1.009 | 0.496 | 0.414 | 1.310 | 0.836 | 0.818 |

Note: $U$ is Theil's $U$-statistic.
[a] $p$-values from parametric bootstrap.
[b] $p$-values from nonparametric bootstrap.

predictive power. The results provide strong evidence that real oil prices play an important role in real exchange rate determination and help predict future real exchange rate returns.

### 6.2. Out-of-sample prediction

Following the seminal paper of Meese and Rogoff (1983), assessing exchange rate models by performing out-of-sample predictions sets a benchmark. We conduct out-of-sample forecasts both on a short horizon ($k=1$) and on long-horizons ($k=16$ and $k=24$). The sample of observations is divided into in-sample and out-of-sample parts. There are 228 in-sample observations (1972:1–1990:12). A recursive scheme of estimation is used. The recursive updating scheme yields 178 one-step-ahead, 162 sixteen-step-ahead, and 154 twenty-four-step-ahead out-of-sample observations. For WTI oil prices, the in-sample period is from 1983:1 to 1987:6 and has 154 observations. This is because the estimating sample begins at 1983:1. Relative forecast accuracy is measured by using Theil's $U$-statistic, which is the ratio of the root-mean-square prediction errors from two competing models. We test the null hypothesis that forecasts based on real oil prices and forecasts based on the naive random walk model are equally accurate ($U=1$). The alternative hypothesis is that forecasts based on real oil prices outperform the random walk forecasts ($U<1$). Tables 10 and 11 report the results for the out-of-sample forecasts. We first focus on those based on the world oil price, the Dubai price, and the Brent price. In the one-step-ahead prediction ($k=1$), all of the

Table 12
Robustness check of model specification

| | World prices | Dubai | Brent | WTI |
|---|---|---|---|---|
| Real oil price ($\beta_1$) | 0.2268 | 0.2255 | 0.2419 | 0.3363 |
| *t*-statistics | 18.4076 | 19.3149 | 19.6963 | 21.814 |
| Real interest rate differential ($\beta_2$) | −0.0005 | −0.0009 | −0.0011 | −0.0124 |
| *t*-statistics | −0.4967 | −0.0469 | −0.6402 | −4.6203 |
| Productivity differential ($\beta_3$) | −0.4379 | −0.3838 | −0.3914 | −0.0245 |
| *t*-statistics | −8.2890 | −7.7376 | −7.0499 | −1.1465 |

Note: The model is $q_{it}=\alpha_i+\beta_1 p_{it}+\beta_2(r_{it}-r^*_{it})+\beta_3(y_{it}-y^*_{it})+\varepsilon_{it}$, where $q_{it}$ is the log of the real exchange rate, $p_{it}$ is the log of the real oil price, $r_{it}-r^*_{it}$ is the real interest rate differential, and $y_{it}-y^*_{it}$ is the log of the productivity differential. Estimates were obtained by using FMOLS.

forecasts based on real oil prices perform better than a random walk: Theil's *U*-statistics are around 0.99, except in the case of the real exchange rate between the Canadian and US dollars. The bootstrapped *p*-values suggest a significantly superior forecasting performance. In general, the longer the forecasting horizon, the smaller are the *U*-statistics. The exceptions are for the real exchange rates between the UK pound and the US dollar and those between the Japanese yen and the US dollar. However, for the WTI price, the evidence is somewhat weaker. In this case, only the one-step-ahead predictions based on real oil prices are superior. Moreover, there is no evidence to support better predictability over longer horizons.

In summary, our evidence suggests that forecasts based on real oil prices perform better than those based on a random walk. Furthermore, predictability is higher over longer horizons. The improvement tends to increase with the forecast horizon. This finding is similar to those of Mark (1995) for nominal exchange rate predictions.

## 7. Robustness

In this section, we provide a robustness check of the model specification. In addition to real oil prices, the two most important determinants of real exchange rate movements from the literature are real interest rate differentials and productivity differentials. The former is motivated by uncovered real interest rate parity (see Meese and Rogoff (1988) and Camarero and Tamarit (2002)) whereas the latter is motivated by the well-known Balassa–Samuelson effect. We consider the following regression model:

$$q_{it} = \alpha_i + \beta_1 p_{it} + \beta_2(r_{it}-r^*_{it}) + \beta_3(y_{it}-y^*_{it}) + \varepsilon_{it}, \tag{8}$$

where $q_{it}$ is the log of the real exchange rate, $p_{it}$ is the log of the real oil price, $r_{it}-r^*_{it}$ is the real interest rate differential, and $y_{it}-y^*_{it}$ is the log of the productivity differential. A rise in $r_{it}-r^*_{it}$ is predicted to appreciate the domestic currency (see Meese and Rogoff (1988)), and an increase in $y_{it}-y^*_{it}$ is expected to appreciate the domestic currency, as predicted by the Balassa–Samuelson model. Thus, we expect $\beta_2<0$ and $\beta_3<0$. Real interest rates are measured by subtracting CPI inflation rates from nominal interest rates. Productivity is measured by subtracting the log of employment from the log of real GDP. Descriptions of the variables used are provided in Table 1.

We report the results from Eq. (8), estimated by FMOLS, in Table 12. Clearly, these results confirm our main finding that a higher real oil price depreciates the real exchange rate in the long run. This suggests that our results are robust. Moreover, the effects of real interest rate differentials and productivity differentials on the real exchange rates are negative, as predicted by the theory.

## 8. Concluding remarks

In this paper, we examined the relationship between real oil prices and real exchange rates for the G7 countries by using monthly panel data from 1972:1 to 2005:10. By using a panel study, we first reexamined whether exchange rates are cointegrated with real oil prices, as has been documented in previous studies, which are primarily time-series case studies of individual countries. We showed that real oil prices may have been the dominant source of real exchange rate movements and that there is a cointegrating relationship between real oil prices and real exchange rates. We used different measures of oil prices: the world price of oil; the United Arab Emirates price of oil (Dubai); the British price of oil (Brent); and the US West Texas Intermediate price of oil. All our empirical results are robust to different measures of oil prices.

Second, along the lines of research on prediction regression tests used to investigate the relationship between exchange rates and fundamentals, we examined the ability of real oil prices to forecast future exchange rate returns. See, for example, Mark (1995), Kilian (1999), Mark and Choi (1997), and Mark and Sul (2001). We assessed the role of real oil prices in predicting real exchange rates over long horizons. Panel predictive regression estimates suggested that real oil prices had significant forecasting power for real exchange rates. Finally, out-of-sample prediction performance using real oil prices outperformed a random walk model, and was better over longer horizons.

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## Appendix A. Supplementary material

Supplementary data associated with this article can be found, in the online version, at doi:10.1016/j.eneco.2006.08.003.

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