Long-run determinants of real exchange rate: New evidence based on panel data unit root and cointegration tests for MENA countries

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Abstract
This paper analyses the main determinants of real exchange rate in the MENA countries. We apply new panel data unit-root tests proposed by Inn, Pesaran and Shin (1997) and new panel data cointegration techniques suggested by Pedroni (1999) by estimating the long-run determinants of real exchange rate and we compare the results with those obtained with conventional time series unit-roots and cointegration tests. Our main finding is that whereas standard time series approach rejects the Balassa-Samuelson hypothesis for 12 countries out of 16, new panel cointegration techniques permit to rescue this hypothesis for MENA countries. Moreover further investigations show that output per capita as well as government consumption, real interest rate differentials and openness degree of the economy also influence real exchange rate.

Keywords: Real Exchange Rate, Economic Growth, Balassa-Samuelson hypothesis, MENA countries, Misalignment, Panel unit-root and cointegration tests.

JEL Classification : E31, F31, C15.
1 Introduction

The relationship between real exchange rate and economic development is certainly an important issue, both from the positive (descriptive) and normative (policy prescription) perspectives. In recent years, policy discussions have included increasing references to real exchange rate stability and correct exchange rate alignment as crucial elements to improve economic performance in emergent countries. Real exchange rate misalignment affects economic activity in developing countries mainly due to the dependence on imported capital goods and specialization in commodity exports. Accessibility to world financial markets which helps to smooth consumption by financing trade imbalance, also plays an important role. Evidence from developing countries is often quoted to support the view that the link between real exchange rate misalignment and economic performance is strong. Cotti and al (1990) argued that in many emergent countries, persistently misaligned exchange rate harmed the development of agriculture, reducing domestic food supply. Indeed, a number of researchers have also pointed out the importance of understanding the main determinants of real exchange rate (Edwards; 1989; Elbadawi and Soto, 1997; Ebadawi, 1994; Ghura and Grennes, 1993).

Evidence from many countries show that successful development results in a currency appreciation with an improvement in the standard of living, whereas a failure in economic development results in a sharp currency depreciation. One of the most important hypotheses with respect to the equilibrium real exchange rate level is the so-called Balassa-Samulson hypothesis (1963) i.e. rapid economic growth is accompanied by real exchange rate appreciation because of differential productivity growth between tradable and non-tradable sectors. However, very few studies (see for example Takatoshi and al (1997)) have been devoted to the analysis of the relationship between the level of economic development and real exchange rate, as was suggested by the seminal paper of Balassa-Samuelson (1964). These relationships permit to determine the equilibrium real exchange rate level to which policy makers refer, to evaluate possible misalignment.

Consequently, one key objective of open economy macroeconomics is to provide policy makers with a framework on equilibrium exchange rate level compatible with economic equilibrium, as well as policy instruments necessary to correct the possible misalignment. So one objective of this paper is to contribute to literature on exchange rate determinants using recent econo-
metric developments and to propose an empirical investigation of the link between real exchange rate and some macroeconomic variables.

Hence we attempt, in a first time, to test empirically the Balassa-Samuelson hypothesis and in second one we extend our analysis to introduce other variables, such as government consumption, real interest rate differential, and trade policy, as the determinants of real exchange rate. Once the determinants of real exchange rate are identified, we estimate the degree of misalignment in real exchange rate and analyze the importance to implement macroeconomic policies to correct misalignment and improve economic performances.

So far much attention has been paid in literature to analyze the real exchange rate determinants using time series econometric techniques. Early cointegration tests such as Engle and Granger (1987) cointegrating regression and Johansen (1988), (1995) maximum likelihood (ML) procedures produce mixed results even if they tend to reject most often the hypothesis of long-run relationship between real exchange rate and some fundamentals for the recent floating exchange rate period. One of the problems of this literature is the low power of such tests against stationarity alternatives in small samples.

Recently methods for non-stationary time series panel, including unit root tests (see for instance Levin and Lin (1993), Quah (1994) Im and al (1996)), and cointegration tests (see for instance Pedroni (1995a, 1997a and 1999) or Blinder, Hsiao and Pesaran (1999)) have been gaining increased acceptance in recent empirical research. The extension of conventional non-stationary methods to panel with both cross section and time series dimensions holds considerable promise for empirical research considering the abundance of available data in this form. Recent applications of these panel tests for cointegration include Canzoneri Cumby and Dida (1996), to productivity and real exchange rate, Obstfeld and Taylor (1996) to international capital mobility, Pedroni (1995b) to endogenous growth theory and Taylor (1996) to historical episodes of purchasing power parity.

In this context, the purpose of this paper is to apply these new panel unit root tests and panel cointegration tests recently developed in the econometric literature to re-analyze empirically the main determinants of real exchange rate. We consider here annual data for 16 MENA countries (Algeria, Bahrain, Egypt, Iraq, Iran, Jordan, Kuwait, Libya, Lebanon, Morocco, Oman, Saudi Arabia, Syria, Tunisia, United Arab Emirates, and Yemen) covering the 1960-1999 period, and we compare the panel data econometric results with those that would have been obtained with conventional unit-root tests and coin-
Integrating techniques. We use In, Pesaran and Shin (1997) panel unit root tests and Pedroni’s panel cointegration methodology (1995a, 1997a and 1999) to analyze the long-run determinants of the real exchange rate. Our main conclusion in that new panel integration and cointegration techniques permit to rescue the original Balassa-Samuelson hypothesis. Moreover, further investigations indicate strong evidence in favor of the importance of output per capita, government consumption, openness, and real interest rate differential as determinants of real exchange rate.

The remainder of this paper is organized as follows. In Section 2 we review major real exchange rate determinants for developing countries. In section 3 we expose first the Balassa-Samuelson benchmark model and then a more general model for real exchange rate determinants in order to give a theoretical support to our econometric analysis. In section 4 we present the panel data unit root tests and panel cointegration methodology that will be used in the empirical application. In section 5 we expose and comment on our econometric results and give policy recommendations for Mena country policy-makers. A final section reviews the main findings.

2 Real exchange rate determinants for developing countries: a survey of literature

Real exchange rate stability and its correct alignment are known to be necessary conditions - though not sufficient - for economic development (Williamson, 1997). This agreement has induced numerous estimations of equilibrium real exchange rates in developing countries during the eighties and early nineties. The first concern of studies on real exchange rate is to find an equilibrium real exchange concept and then to estimate its long run level. Indeed once this has been done one can determine the necessary adjustments to reach equilibrium. Moreover equilibrium real exchange rate determinant identification enables to forecast its evolution and then to choose the appropriate measures to remedy possible divergences and to determine the necessary adjustments with regards to economic policy purposes.

Real exchange rate determinants and the effects of misalignment have been the subject of numerous theoretical and empirical works. However in the optics of equilibrium real exchange rate evaluation for developing countries, the Edward model (1988) is the most extensively used framework. So
Equilibrium real exchange rate is defined in the Edwards model as the relative price of traded and non-traded goods which insures simultaneously the internal and external balances of the economy with underlying capital flows. Long term real exchange rate depends exclusively on real variables, called by the author fundamentals: the terms of trade, public expenditures, the openness of the economy, technical progress and capital flows. Therefore, equilibrium exchange rate is not a constant, on the contrary it evolves following a trajectory determined by fundamental modifications. In the short run, real exchange rate proceeds of real and monetary variable combinations. The vector error correction model including variables expressed in difference in the short run, allows to describe the real exchange rate dynamics towards its long term target. Edwards (1989) also investigates the relationship between real exchange rate misalignment and economic performances and concludes that real exchange rate difference with regards to its equilibrium level has a negative effect. Edwards’ works have influenced several other authors on real exchange rate determinants and misalignment effects. Cotti and al (1990) for instance confirm that for some Latin American countries real exchange rate instability has handicapped exportation growth, whereas Asian exportation growth was for the most part accounted for by real exchange rate stability. Furthermore they indicate that the weak performance of African agriculture and economic activity in general are partly due to real exchange rate misalignment. Other works by Sekkat and Varoudakis (1998) strengthen the idea according to which the chronic misalignments of real exchange rate are a major factor of the weak economic performances of developing countries. Ghura and Grennes (1993) show on a panel of African countries that real exchange rate misalignment negatively affects economic growth, exports, investment and saving. In the works of Ghura and Grennes (1993), Aron and al (1997), Cotti and al (1990) Elbadawi and Soto (1997) taken as a whole, real exchange rate determinants are mainly the terms of trade, the openness degree of the economy, imports and capital flows. The initial works on real exchange rate determinants have used standard univariate econometric estimation methods whereas more recent ones have used cointegration techniques developed by Engle and Granger (1987) and Johansen ((1988), (1995)) which permit to test the existence of a long-run relationship.
between the variables under study. Such investigations include the works of Bayés and al (1999), Elbadawi and Sote (1997), Feyzioglu (1997), Kedenge (1998) and Faruque (1997). On the contrary panel data cointegration methods are still very rare and among studies that used such methods, one finds MacDonald and Nagaysau (2000) who have estimated a long run relationship between exchange rate and interest rate differential and also several studies on purchasing power parity (see for instance Pedroni’s (1995b)).

**GDP as an equilibrium real exchange rate determinant**

The so-called Balassa-Samulson hypothesis con..rms that during development process, productivity increases in the tradable sector tend to be higher than those in the non-tradable one. It follows, if one supposes that wage levels are equal between the two sectors (because of the inter-sectorial mobility of work) a relative increase of the relative prices of non-traded goods. So, a key nding of the original Balassa paper is that there is a positive relationship between aggregate output per head and real exchange rate. In addition, Balassa was the rst author to test formally the proposition that richer countries have a higher real exchange rate. He regressed real exchange rate on GDP per capita and found that GDP per capita growth is positively correlated with real exchange rate appreciation. Balassa (1973) presents similar results ndings that richer countries have a higher price level. The following studies deal with the cross-section implication of Balassa-Samuelson. Hseih (1982) was the rst to look at time-series implications. His study focused on Japan’s and Germany’s exchange rates vis a vis the United States for the 1954-1974 period. He found that productivity differential variables were signi..cantly and of the correct sign. Marston(87) and Edison and Klovland (1987) also present evidence for the Balassa-Samuelson effect. The evidence from later studies is somewhat mixed. Froot and Rogo (1991) show that the correlation between productivity and real exchange rate is weak at best. Asea and Mendoza (94), take a di..erent approach and conclude that the productivity differential between traded and non-traded goods are extremely signi..cant to explain changes in the relative prices of non-traded goods within each country.

The openness degree of the economy as an equilibrium real exchange rate determinant
The openness degree of the economy is an equilibrium real exchange rate determinant since an increasing openness towards foreign countries, leads to an import price decrease and so of the price of tradable goods price, by reducing customs tariffs on imports for instance. A tradable goods decrease induces a real exchange rate appreciation. On the contrary a customs tariffs rise on imports has an opposite effect. However the increase of import demand which follows leads to a rise of foreign currency demand and involves consequently a real exchange rate appreciation. Thus the effect of trade liberalization on real exchange rate is a priori indeterminate. Fazes and al (1999) show that for the Ivory Coast and Burkina Faso, trade liberalization has induced a real exchange rate depreciation.

Public expenditures as an equilibrium real exchange rate determinant

The impact of public expenditures on real exchange rate both depends on its level and on its distribution between tradable and non-tradable goods. Indeed if the horizon is composed of two periods and if one supposes that in the first one the government increases its non-tradable goods consumption and that its expenditures are financed by borrowing on domestic and international markets, the demand increase induces a non-tradable goods price rise and so a real exchange rate appreciation. However for the second period the government will need to increase taxes to finance these debts and so household available income will decrease and it is the same for global demand. A non-tradable goods demand decrease leads to a price decrease and then to a real exchange rate depreciation. Thus, in that case the effect of a public expenditure rise on real exchange rate is a priori indeterminate. Edwards (1989) finds that a public expenditure increase induces a real exchange rate appreciation for a set of 12 developing countries.

Real interest rate differential as an equilibrium real exchange rate determinant

The modernization and privatization effort in developing countries is often accompanied by a capital movement liberalization and a financial market development. Capital flows which follow these liberalization efforts will have significant effects on real exchange rate and one of the major challenge for policy decision-makers is how to limit the effects of these flows on real
exchange rate and trade balance. Agénor and Hořašmaier (1998) confirm that capital flows have led to a real exchange rate appreciation for Latin American countries.

3 Real exchange rate determination: a theoretical framework

3.1 The Balassa-Samuelson hypothesis benchmark model

In order to explicit the nature of the relation to be tested on panel data in section 4, we present here the Balassa-Samuelson hypothesis as a minimal theoretical framework for real exchange rate determination.

For this purpose let us consider a small open economy composed of a set of homogeneous firms. The representative firm produces two goods: a tradable commodity on the world market and a non-tradable one for domestic demand. It is supposed besides that tradable goods production requires both capital and labour, whereas non-tradable goods production only uses labour. The competition is supposed to be perfect and it ensures that production factors are paid at their marginal productivity; labour factor mobility ensures equal pay. Labour supply is supposed to be constant and all variables are expressed in terms of tradable goods.

As noted by Rogoξ and Obstfeld (1996), in the absence of nominal rigidity, equilibrium real exchange rate will only depend on αer. Thus in what follows we will present a partial equilibrium model where the demand side will be absent.

3.2 Firms’ behaviours

The representative firm maximizes its intertemporal profit expressed in terms of tradable goods under its constraints of technology and capital accumulation, that is:

\[
\max_{0}^{Z} (y_{e}(k; l_{e}) + p_{yn}(l_{n}) \cdot \omega_{l} \cdot i) e^{i \cdot r \cdot t} \cdot dt
\]

where

\[
sc k \equiv i \cdot \omega_{k}
\]
where,

\( y_e \) denotes the production of tradable goods;

\( y_n \) denotes the production of non-tradable goods;

\( p \) denotes the relative price of non-tradable goods in terms of tradable ones;

\( i \) denotes investment

\( w \) denotes wages;

\( k \) denotes capital;

\( l = l_n + l_e \) is labour supply.

### 3.3 Equilibrium

The equilibrium is defined as follows

\[
\frac{dy_e}{dk} = r
\]

\[
\frac{pdy_n}{dl_n} = \frac{dy_e}{dl_e} = w
\]

\[
\frac{dy_e}{dl_e} = \frac{dy_n}{dl_n} = 1
\]

We thus obtained the following relationships between relative price (tcr) and productivity ratio:

\[
\frac{dy_e}{dy_n} = p
\]

For Cobb-Douglas functions, this relation expresses as:

\[
p = \frac{\theta y_e}{l_n}
\]
where \( \bar{\alpha} \) and \( \bar{\beta} \) are the production-labour elasticities respectively for tradable and non-tradable sectors and \( \mu_n, \mu_e \) the labour average productions for the two sectors.

These relationships expressed well the Balassa Samuelson hypothesis and indicate that relative price is a function of the productivity ratio of the two goods. Thus a faster increase of tradable goods productivity than of non-tradable ones leads to a real exchange rate appreciation. We now need to determine a relationship between per capita GDP (that will be an indicator of development) and exchange rate.

In fact, the production of the economy, expressed in terms of tradable goods writes as:

\[
y = y_e + p y_n
\]

that is for the Cobb-Douglass case:

\[
y = \alpha \mu_e l_e + \beta \mu_n l_n
\]

where \( \bar{\alpha} \) and \( \bar{\beta} \) are respectively the labour elasticities in the tradable and non-tradable sectors.

We consider now the following simplification \( \bar{\alpha} = \bar{\beta} \); which enables us to obtain the below relationship:

\[
\frac{y}{l} = p \mu_n
\]

According to this equation for a given level of non-tradable productivity, relative price is all the higher as per capita production is high.

Furthermore real exchange rate quoted to certain is defined as:

\[
e = \frac{P}{EP}
\]

where,

\( E \) denotes nominal exchange rate,
\( P \) denotes general domestic price index,
\( P^* \) denotes general foreign price index.

If we suppose that the consumer’s basket contains two commodities, we can express the general price index as:
\[ P = P_e P_n^1 e P^* = (P_e P_n^1)^1 \]

Then, following Balassa and Samuelson and if we suppose that parity purchasing power in the tradable sector is verified, we will have:

\[ \log(e) = (1 - \eta) \log(P) - (1 - \eta) \log(P^n) \]

and taking the above analysis into account, we obtain:

\[ \log(e) = i (1 - \eta) \mu + (1 - \eta) \log \left( \frac{Y}{P} \right) i \log(P^n) \]

These relationships indicate that the development level of an economy determines its long-term real exchange rate behaviour. The nature of non-tradable goods makes its productivity vary independently of foreign exchanges. Furthermore as shown by Roldos (1995), this sector is weakly intensive in capital and uses rather non reproducible factors. This independence entails that in the long term productivity gain evolution can be supposed to be exogenous.

### 3.4 A more general framework

In this section we expose an extension of the benchmark model where the equilibrium real exchange rate is as a path upon which an economy maintains both internal and external balances. The equilibrium real exchange rate depends not only on productivities but also on some other real variables. The model used is developed by Montiel (cited by Mkenda 2001) and illustrate the importance of government consumption, real interest rate as the determinants of equilibrium real exchange rate. Real exchange is defined as the relative price of traded and non-traded goods. That is:

\[ \text{R E R} = e = E \frac{P_T^n}{P_N^n}, \]

where, \( P_T^n \) is the world price for traded goods and \( P_N^n \) is the price of non-traded ones, and \( E \) is the nominal exchange rate.
The equilibrium real exchange rate is such as both internal and external balances are attained. Internal equilibrium is attained when the market of non-traded goods clears:

\[ Y_N(e) = (1 - \% C) + G_N \]

where,
- \( Y_N \) is the production of non-traded goods;
- \( G_N \) is the government consumption of non-traded goods;
- \( \% \) is the share of traded goods in total consumption; and
- \( C \) is total private consumption.

The internal balance enables us to extract a negative relationship between real exchange rate and private consumption. In fact, a rise in private consumption causes an excess of demand for traded as well as non-traded goods and so an increase in the relative prices of non-traded ones (the price of traded goods depends on the world market). The appreciation of the real exchange rate, caused by an increase of the price of non-traded goods, will switch supply towards non-tradeable goods, and demand towards tradable ones.

External balance is defined by the following equation of the current account balance:

\[ B = rB + Y_T(e) - G_T - \% C \]

where,
- \( B \) is net foreign assets;
- \( Y_T \) is the domestic supply of traded goods;
- \( G_T \) is government spending on traded goods;
- \( r \) is the world interest rate.

From the external balance, \( \dot{B} = 0 \); we obtain a relationship between private consumption and real exchange rate. A rise in private spending generates a current account deficit and in order to restore equilibrium, the real exchange rate must depreciate. The depreciation of real exchange rate switches demand towards non-tradeable goods, and supply towards tradable ones.

The equilibrium real exchange rate is such as both internal and external balances are achieved and defined as:

\[ e^* = e(G_N; G_T; B; r) \]
Indeed, if we distinguish two kinds of tradable goods (exported and imported ones) whose prices are determined on the world market, and if before liberalization tariff measures exist on imports, real exchange rate will be defined as follows:

\[ e = \frac{E[(1 + t)P_m + (1 - \bar{t})P_x]}{P_N} \]

where \( P_m \) and \( P_x \) are world prices of imported and exported goods and \( t \) is the tariff on imports. Trade liberalization will lead to a progressive decrease of \( t \) and a real exchange rate appreciation.

In our empirical model we try to take these different variables into consideration and as an extension of this model we include an indicator of capital inflows. As unfortunately we don't have data on public expenditure variation composition we approximate it by the share of total government consumption in GDP. The capital inflows will be measured by the real interest rate differential between domestic and foreign economies. Productivity will be approximated by the per capita GDP level and the ratio between the total of exports and imports and GDP will be considered as an indicator of the openness degree of the economy.

4 Econometric methodology

Given the previous theoretical analysis, the two cointegration relationships to be tested can be written as:

\[ \log(e_t) = \bar{r} + \gamma t + \phi \log(y_t) + \delta_t \]

In this relation, the variable trend \( t \) reflects the non-tradable sector productivity evolution and we expect \( \gamma \) to be positive. We also expect \( \phi \) to be positive since an increase of real exchange rate implies an appreciation. As far as the sign of the constant is concerned it is a priori negative and a rise in tradable goods prices leads to an exchange rate depreciation. The small economy hypothesis means that foreign price is totally exogenous.
for the more general model:

$$
\log(e_t) = \beta_0 + \beta_1 \log(y_t) + \beta_2 \log(\text{opness}) + \beta_3 \log(i-i^*) + \beta_4 \log(G_t) + \epsilon_t
$$

openness is the indicator of trade policy and the sign of $\beta_2$ is a priori undefined, $i-i^*$ is the real interest rate differential and we expect $\beta_3$ to be positive, and in the same way the sign of $\beta_4$ is a priori positive.

We now present the panel unit root tests and panel cointegration tests that we will use in the empirical application reported in section 4.

### 4.1 Panel unit root tests

Initial methodological work on non-stationary panels focused on testing unit roots in univariate panels. Quah (1994) derived standard normal asymptotic distributions for testing unit roots in homogeneous panels as both time series and cross sectional dimension grow large. Levin and Lin (1993) derived distributions under more general conditions that allow for heterogeneous fixed effects and time trend. More recently, Im, Pesaran and Shin (1997), studied the small properties of unit root tests in panels with heterogeneous dynamics and proposed alternative tests based on the mean of individual unit-root statistics. In this paper we shall apply Im and al (1997) unit-root test (called IPS after) since it is more powerful than those of Quah (1994) and Levin and Lin (1993) used in existing studies.

Levin and Lin’s test is considered more general than those of Quah since it explicitly takes heterogeneity and correlation between units into account. However as shown by Papell (1997) it suffers from size distortion without being able to correct adequately serial correlation. Using Monte Carlo simulations, he showed that the finite sample critical values are greater than those in Levin and Lin (1993). For quarterly data, the critical values are 11% higher (on average) than those reported by Levin and Lin and for monthly data, they are 3% higher.

The test proposed by Im and al (1997) permits to solve Levin and Lin’s serial correlation problem by assuming heterogeneity between units in a dynamic panel framework. Furthermore as shown by Im and al via Monte Carlo simulations it has higher power than that of Levin and Lin. IPS (1997) proposes two statistics: a Maximum Likelihood Statistics, called Lbar, and a Student statistic $t_b$. These two statistics are based on individual Augmented
Dickey-Fuller (ADF) regressions. Since an appropriate ADF regression will correct the serial correlation in the data, the IPF panel unit-root test takes care of serial correlation automatically. In our empirical work of section 4 we shall use the $t_b$ statistic instead of the $Lbar$ one since IPS's Monte Carlo experiments have shown that it is the more powerful even for a value of $N$ inferior to 5. This statistic can be expressed as:

$$t_b = \frac{p}{N} \left( \sum_{i=1}^{N} (t_{iT}) \right) \frac{E(t_t)}{\text{Var}(t_T)}$$

where $t_{NT} = \frac{1}{N} \sum_{i=1}^{N} t_{iT}$ is an average of the $t$ individual student statistic in a conventional time series unit-root analysis, $E(t_t)$ and $\text{Var}(t_T)$ are respectively the mean and variance of $t_{iT}$ under the null hypothesis that the series are integrated of order one with $N N_! 1$:

IPS shows that under the null hypothesis of non-stationarity, the $t_b$ statistic follows the standard normal distribution asymptotically.

4.2 Panel cointegration tests

In this paper we shall apply Pedroni's cointegration test methodology (1995a, 1997a and 1999) to analyse the Balassa-Samuelson hypothesis. Pedroni (1995) studied the properties of spurious regressions and tests for cointegration in heterogeneous panels and derived appropriate distributions for these cases. These allow us to test the presence of long run equilibria in multivariate panels while permitting the dynamic and even the long run cointegrating vectors to be heterogeneous through individual members. Like the IPS panel unit-root test, the panel cointegration tests proposed by Pedroni also take heterogeneity into account by using specific parameters which of course are allowed to vary across individual members of the sample. Pedroni (1997a and 1999) derived the asymptotic distributions and explored the small sample performances of seven different statistics to test panel data cointegration. Out of these seven statistics, four are based on pooling, what is often referred to as the Within dimension (called "panel" after), and the last three are based on the Between dimension (called "group" after). These different statistics are based on a model that assumes that cointegration relationships are heterogeneous between individual members and are defined as:

For the Within statistics
For the Between statistics

\[ Z_{1/2}^B = \left( \prod_{i=1}^{N} \prod_{t=1}^{T} L_{i1}^{2} \left( \mathbf{e}_{i}^{2} \mathbf{e}_{i} \right) \right)^{1/2} \left( \mathbf{e}_{i1} \mathbf{e}_{i1} \right)^{1/2} \mathbf{e}_{i1}^{2} \mathbf{e}_{i1}^{2} : \text{Rho}_{\text{stat}} \]

\[ Z_{1/2}^w = \left( \mathbf{e}_{N1}^{2} \mathbf{e}_{N1} \right)^{1/2} \left( \mathbf{e}_{11} \mathbf{e}_{11} \right)^{1/2} \mathbf{e}_{11}^{2} \mathbf{e}_{11}^{2} : \text{PP}_{\text{stat}} \]

\[ Z_{1/2}^{pp} = \left( \mathbf{e}_{N1}^{2} \mathbf{e}_{N1} \right)^{1/2} \mathbf{e}_{11}^{2} \mathbf{e}_{11}^{2} : \text{Adf}_{\text{stat}} \]

\[ Z_{1/2}^{w} = \left( \mathbf{e}_{N1}^{2} \mathbf{e}_{N1} \right)^{1/2} \mathbf{e}_{11}^{2} \mathbf{e}_{11}^{2} : \text{v}_{\text{stat}} \]

For the Between statistics

\[ Z_{1/2}^B = \left( \mathbf{e}_{i1} \mathbf{e}_{i1} \right)^{1/2} \left( \mathbf{e}_{i1} \mathbf{e}_{i1} \right)^{1/2} \mathbf{e}_{i1}^{2} \mathbf{e}_{i1}^{2} : \text{Rho}_{\text{stat}} \]

\[ Z_{1/2}^{B} = \left( \mathbf{e}_{i1} \mathbf{e}_{i1} \right)^{1/2} \left( \mathbf{e}_{i1} \mathbf{e}_{i1} \right)^{1/2} \mathbf{e}_{i1}^{2} \mathbf{e}_{i1}^{2} : \text{Adf}_{\text{stat}} \]

\[ Z_{1/2}^{pp} = \left( \mathbf{e}_{i1} \mathbf{e}_{i1} \right)^{1/2} \mathbf{e}_{i1}^{2} \mathbf{e}_{i1}^{2} : \text{pp}_{\text{stat}} \]

with,

\[ b = \frac{1}{r} \prod_{s=1}^{r} \left( \frac{1}{t} \sum_{t=s+1}^{T} \mathbf{b}_{it} \mathbf{b}_{it} \right) \mathbf{b}_{it} \mathbf{b}_{it} \mathbf{s}_{i1}^{2} \]

\[ b_{2}^{2} = \frac{1}{r} \prod_{t=s+1}^{T} \mathbf{b}_{it}^{2} ; \mathbf{b}_{2}^{2} = \mathbf{s}_{i1}^{2} + 2 \mathbf{b}_{it} \mathbf{b}_{it} \]

\[ \mathbf{b}_{2}^{2} = \frac{2}{r} \mathbf{b}_{it}^{2} ; \mathbf{b}_{2}^{2} = \mathbf{s}_{i1}^{2} + 2 \mathbf{b}_{it} \mathbf{b}_{it} \]

\[ \mathbf{b}_{2}^{2} = \frac{1}{r} \prod_{t=s+1}^{T} \mathbf{b}_{it}^{2} ; \mathbf{b}_{2}^{2} = \frac{1}{r} \prod_{t=s+1}^{T} \mathbf{b}_{it}^{2} ; \mathbf{b}_{2}^{2} = \frac{1}{r} \prod_{t=s+1}^{T} \mathbf{b}_{it}^{2} ; \mathbf{s}_{i1}^{2} \]
and where the residuals are extracted from the above regressions:

\[ \mathbf{a}_t = \mathbf{b}_t \mathbf{X}_{it} + \mathbf{e}_t; \]

\[ \mathbf{a}_t = \mathbf{b}_t \mathbf{X}_{it} + \sum_{k=1}^{p} \mathbf{b}_k \mathbf{C}_k \mathbf{a}_{t-k} + \mathbf{e}_t; \]

\[ \mathbf{y}_t = \mathbf{M}_t \mathbf{X}_{mt} + \mathbf{b}_t; \]

Note that in the above writings \( L_i \) represents the \( i^{th} \) component of the Cholesky decomposition of the residual Variance-Covariance matrix, \( \mathbf{b} \) and \( \mathbf{e} \) are two parameters used to adjust the autocorrelation in the model, \( \mathbf{Y}_i \) and \( \mathbf{s}_i^2 \) are the contemporaneous and long-run individual variances.

Pedroni has shown that the asymptotic distribution of these seven statistics can be expressed as:

\[ \hat{A}_{NT} = \mathbf{b} \mathbf{p} \mathbf{N} \mathbf{V}^{-1} \mathbf{N} (0; 1) \]

where \( \hat{A}_{NT} \) is the statistic under consideration among the seven proposed, \( N \) and \( T \) are the sample parameter values and \( \mathbf{b} \) and \( \mathbf{p} \) are parameters tabulated in Pedroni (97).

In terms of power Pedroni (97) showed that for values of \( T \) larger than 100, all the proposed seven statistics BGVD do fairly well and are quite stable. However for smaller samples (\( T \) inferior to 20) the \( Z_{\mathbf{b}}^p \) statistic is the most powerful, followed by the \( Z_{\mathbf{w}}^y \) and \( Z_{\mathbf{w}}^t \) ones. The finite sample distribution for the seven statistics have been tabulated by Pedroni (1997a) via Monte Carlo simulations. The calculated test statistics must be smaller than the tabulated critical value to reject the null hypothesis of absence of cointegration.

5 Empirical results

We present in this section a study of panel data unit root tests and cointegration tests using data:

- over a panel of 16 Mena Countries (Algeria, Bahrain, Egypt, Iraq, Iran, Jordan, Kuwait, Libya, Lebanon, Morocco, Oman, Saudi Arabia, Syria, Tunisia, United Arab Emirates, and Yemen) over a 39 year period (the estimation sample covers the 1960-1998 period) for the minimal Balassa-Samuelson framework,
- and over a panel of 7 MENA countries (Bahrain, Egypt, Jordan, Kuwait, Morocco, Syria, and Tunisia) over a 20 year period (the estimation sample covers the 1978-1998 period) for the more general theoretical model.

All the data are taken from the CHELEM French database of the CEPII and World Bank Data Base. They are all annual time series observation. Using panel techniques, we test:

1. in a .rst step, according to the benchmark model (presented in the previous section) the long-run relationship between real exchange rate and GDP per capita (as indicator of development) as described below:

   \[
   \log(RER_{it}) = \beta_0 + \beta_1 t + \beta_2 \log(Y_{it}) + \beta_3 \log(P_{it}) + \beta_4 \log(Y_{it}) + \gamma_{it} \]

   where \( RER \) is the real exchange rate defined as the ratio between the domestic price index and the foreign price index with respect to the USA deflated by the nominal exchange rate (so an increase of \( RER \) indicates an appreciation).

2. and in a second one, according to the more general model the long-run relationship between real exchange rate, GDP per capita, the openness degree of the economy (ouv), the real interest rate differential (idi), and public expenditures (G) as described below:

   \[
   \log(RER_{it}) = \beta_0 + \beta_1 \log(Y_{it}) + \beta_2 \log(A_{it}) + \beta_3 \log(idd_{it}) + \beta_4 \log(G_{it}) + \gamma_{it} \]

   where \( G \) is the ratio between global public expenditures and the GDP level, ouv is the ratio between exports and imports on the one side and the GDP on the other, and \( i \) is the differential between domestic and foreign real interest rates.

5.1 Unit-Root test results

We shall report in this sub-section the results of two kinds of unit-root tests: the conventional time series ones and the In and al (1997) panel data ones. We consider two sample periods and annual data (1960-1998 for real exchange rate and GDP per capita, 1976-1998 for government consumption, openness of the economy, and real interest rate differential).
The analysis first step is simply to look at the data univariate properties and to determine their integratedness degree. Theoretically a process is either I(0), I(1) or I(2). Nevertheless in practice many variables or variable combinations are borderline cases, so that distinguishing between a strongly autoregressive I(0) or I(1) process (interest rates are a typical example), between a strongly autoregressive I(1) or I(2) process (nominal prices are a typical example) is far from being easy. We have therefore applied a sequence of standard time series unit root tests (Schmidt and Phillips test (1992), Kwiatkowsky, Phillips and Shin test (KPSS) (1992) and the efficient unit-root tests suggested by Elliott et al. (1996) (which we shall refer hereafter as the ERS test)), to investigate which of the I(0), I(1), I(2) assumption is most likely to hold. The results of these conventional unit-root tests are reported in table 1 in the appendix. They indicate that the unit-root null hypothesis cannot be rejected at the 5% level for the five variables under consideration (RER,G,GDP,OUV,i-i*) for most of all our 16 countries. The only exception is for Algeria, Bahrain, Egypt, Libya, and Yemen where the KPSS tests indicate that RER and/or per capita GDP are stationary around a linear trend. However the Schmidt-Phillips and Elliott tests confirm the existence of a unit-root in these series. We have also applied those three tests on the variables taken in first difference and we find evidence in favor of the rejection of the non-stationary hypothesis in RER, per capita GDP, G, OUV and i-i* for all countries, which led us to conclude that our series are well characterized as an I(1) process, some with non-zero drift for some countries.

As far as the IPS (1997) panel data unit-root test is concerned (which we have applied for a model with a constant, and for both a constant and a trend) it indicates that for all 5 variables of all Mena countries under study as well as for the two sample periods, the unit-root hypothesis cannot be rejected (see tables 2a and 2b in appendix).

5.2 Cointegration test results

The results of our cointegration analysis are reported in the appendix and as for the unit-root tests presented in the previous sub-section we also consider both time series cointegration tests (see table 3 in the appendix) as well as panel cointegration tests suggested by Pedroni (see table 4). Table 3 reports the results of Johansen (1988), (1995)) conventional time series cointegration tests and they indicate that for 5 countries out of 16 (Kuwait, Libya, Saudi Arabia, Syria, Tunisia) the hypothesis of absence of cointegration cannot
be rejected at a 5% level of significance. Thus, on the basis of cointegration time series tests, we find strong evidence in favor of the rejection of the Balassa-Samuelson hypothesis, since it appears difficult to put in evidence a cointegrating relationship statistically significant between real exchange rate and Per capita GDP. Note that these tests have been implemented only for the Benchmark model and not for the more general one because in the latter case we only have twenty years on annual data, which is not enough for the robustness of cointegration tests.

However panel data cointegration tests enable us to rescue the Balassa-Samuelson hypothesis for MENA countries, since we were able to obtain more robust results. In effect, the results reported in table 4a clearly indicate that for all the 16 countries considered the hypothesis of the existence of a long-run relationship between real exchange rate and Per capita GDP is largely confirmed at a 5% level of significance. In addition almost all coefficients have the expected sign (see table 5a). Concerning GDP we find however counter-intuitive results (i.e. a negative impact on real exchange rate) for 4 countries (Kuwait, Oman, Saudi Arabia, United Arabic Emirates). Nevertheless, the theoretical long-run relationships between real exchange rate and Per capita GDP is strongly confirmed with the panel data approach.

Moreover for the more general theoretical model we also find significant long-run relationships between real exchange rate and its determinants (real interest rate differential, GDP, government consumption, public expenditure) for the 7 countries under consideration (see table 4b). For GDP we find an opposite expected sign (that is a negative impact on real exchange rate) for 2 countries (Bahrein and Jordan) (see table 5b). As far as public government expenditure is concerned the results are mixed: for Bahrain, Kuwait and Morocco, these expenditures affect negatively real exchange rate, which implies that an increase in government consumption causes a real exchange rate depreciation. For the 4 other countries (Egypt, Jordan, Syria, and Tunisia) real exchange rate and public government expenditures are positively correlated. For the openness degree of the economy we find a positive relationship for 3 countries out of 7 (Bahrein, Egypt, Kuwait), which means that an increase of the liberalization of the economy have induced a real exchange rate appreciation. Finally for real interest differential we find counter-intuitive results for three countries (Egypt, Kuwait and Tunisia), which implies that for these countries an increase of the real interest rate differential has led to a real exchange rate depreciation.
All these results show the superiority of the panel data cointegration tests, which are more powerful than conventional time series ones and underline the necessity to be cautious when interpreting the results of usual time series tests for samples of relatively moderate size.

5.3 Misalignment assessment

As we have already mentioned it before the interest to analyze equilibrium real exchange rate determinants is to estimate the misalignment degree. In this sub-section we try to assess the misalignment degree of the different currencies of our sample to question the relevance of the undertaken efforts to improve competitiveness. To estimate the misalignment degree we proceed as follows: the equilibrium real exchange rate level will be defined as the trend part of the estimated real exchange rate cointegrating relationships. Then, to extract the trend we use the Hoddrick-Prescott (HP) filter. Thus the misalignment degree will be defined as the relative difference between effective real exchange rate and the equilibrium level. The misalignment indexes reported in graph 1 in the appendix permit to distinguish three periods for the set of 16 MENA countries under study except for Syria, Yemen and Lebanon. The first period ends up in the early seventies in 1973, date of the rst oil crisis and it is characterized by strong fluctuations. Real exchange rate strong instability during this period can be explained by the fact that for all the countries of our sample, it is the reconstruction period and the nature of economies was not still very clear. The second period begins in 1974 for the majority of the MENA countries except for Bahrain and Egypt where it begins a bit later, in the early eighties. This period is characterized by a real exchange rate appreciation which can be explained for the most part by the oil price rise. Indeed, the majority of these countries are oil exporters and the rise of oil costs has induced currency revenue which have led to a demand increase of domestic currency and so to a real exchange rate appreciation. The oil crisis backlash as well as the gulf war will put an end to this real exchange rate appreciation period, but the effect appears with some delay in the late eighties and the early nineties for some countries. For Tunisia, Morocco, Jordan and Egypt the depreciation is rather caused by voluntarist policies which have accompanied the structural adjustment programs whose goal was to re-establish competitiveness on the international market. However it must be mentioned that real exchange rate evolution still remains dependent on
the international economic situation which results in a high sensibility to exogenous shocks. Policy makers find themselves then compelled to use a set of economic political measures to face the possible fluctuation of real exchange rate. Thus they have to choose between adopting a passive policy which consists in intervening by adapted measures to stabilize the exchange rate when the economy faces external shocks and so avoid their consequences on activity, or to strengthen the independence from abroad, which implies to diversify more the economy and develop substitution industries to replace imports. For Syria, Yemen and Lebanon real exchange rate appears relatively more unstable and follows trajectories different from the other countries of the sample. Economic measures are then possible to stabilize real exchange rate and reduce the effects on activity.

6 Conclusion

In this paper we have re-examined empirically the real exchange rate determinants for 16 MENA countries using new panel unit root tests proposed by In, Pesaran and Shin (1997) and panel cointegration tests recently developed in the econometric literature by Pedroni’s (1995a, 1997a and 1999), which permit to get round the problem of insufficiency of available data. We have also compared our results with those that would have been obtained with conventional time series unit root tests (as Schmidt-Phillips (1992) tests, KPSS (1992) test and the efficient unit-root tests suggested by Elliott et al. (1996)), as well as conventional cointegration tests (as Johansen (1988), (1995) tests).

Our main finding is that whereas time series tests appear to be unable to put in evidence the existence a long-run relationship between real exchange rate and per capita GDP for more than two third of MENA countries (the hypothesis of cointegration was rejected for 11 countries out of 16), which implies the empirical rejection of the Balassa-Samuelson hypothesis, the recourse to new panel data unit-root tests and cointegration techniques permit to rescue this hypothesis for all MENA countries. Moreover, most of the estimated coefficients have the expected sign and confirm that economic development is accompanied by a real exchange rate appreciation. This result is clearly in favour of the Balassa-Samuelson hypothesis and advocates the taking of this effect into account for the analysis of long-run exchange rate
behaviour. Furthermore we show that other variables than GDP per capita such as real interest rate differential, government consumption, the degree of openness of the economy has also a significant effect on the equilibrium real exchange rate for 7 countries (Bahrain, Egypt, Jordan, Kuwait, Morocco, Syria, and Tunisia). Besides, the analysis of real exchange rate misalignment shows three different periods in real exchange rate fluctuations. We conclude that real exchange rate is highly affected by exogenous shocks. In fact the first oil crisis has caused real exchange rate appreciation in the oil producing countries and the oil crisis backlash decrease in oil prices will put an end to this appreciation and real exchange rate has continued to depreciate since 1986. We also note that countries like Tunisia and Morocco who had succeeded in their adjustment programs have the least degree of misalignment. Real exchange rate stabilization efforts are thus necessary to improve economic performances and export; diversification is needed to protect the economy from exogenous shocks.

7 References


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