# Differences in Long Run Exchange Rate Pass-Through Into Import Prices In Developing Countries: An Empirical Investigation

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**Abstract**

This paper investigates the exchange rate pass-through into import prices in a sample of 24 developing countries over the period from 1980 to 2003. We estimate a pass-through equation determined by a combination of the nominal exchange rate, the price of the competing products, the exporter's costs and demand conditions. We adopt non-stationary panel estimation techniques and tests for cointegration. In the long run, homogeneity of pass-through rates across countries can be rejected. Moreover, we show that most of these differences in exchange rate pass-through into import prices are due to three macroeconomics determinants: exchange rate regimes, trade barriers and inflation regimes.

**JEL classification**: F14, C13, C23

**Keywords**: Exchange rate pass-through; Developing countries; Panel cointegration; Non-stationary panel estimation

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# 1 Introduction

The question of how exchange rate changes affect the price level is once again a popular question of research. Typically, this process is called exchange rate pass-through, referring to the degree to which exchange rate is passed through to price level changes. Since the 1980s, there has been revived and increasing interest in exchange rate pass-through into import prices (defined as the percentage change in import prices expressed in domestic currency caused by one percent change in exchange rate). There has been a large number of empirical studies on exchange rate pass-through. These studies can be divided into three categories. The first category has focused on examining exchange rate pass-through into disaggregated import prices of specific domestic industries (see Bache, 2002 and Goldberg, 1995). The second one has examined exchange rate pass-through into aggregate import prices (Hooper and Mann, 1989; Campa and Goldberg, 2002; Webber, 1999). The third category has analyzed exchange rate pass-through into consumer price index (CPI) (Bailliu and Fujii, 2004; McCarthy, 2000; Choudri, Faruquee and Hakura, 2003). The growing research on exchange rate pass-through at the industry-specific and aggregate level is partly motivated by the rise in the industrial organization and strategic trade theory. On the other hand, the empirical studies on the exchange rate pass-through into CPI grow along with development in the new open economy macroeconomic models.

Most of the empirical studies on exchange rate pass-through using different empirical methodologies have focussed on the industrialized countries, in particular the United States and Japan. Menon (1995) surveyed 48 studies on the exchange rate pass-through. He found that most of the research in this area is done on U.S and Japan data. Goldberg and Knetter (1997) noted that, in the 1980s, research on exchange rate pass-through is dominated by the analysis of pass-through to the U.S. However, a few studies on exchange rate pass-through have been done for developing countries (see Alba and Papell (1998), Anaya (2000) and Garcia and Restrepo (2001)).

The first aim of our paper is to contribute to the empirical analysis of exchange rate pass-through into import prices in some developing countries. In particular, we implement an empirical analysis that makes use of the new panel data cointegration techniques. The objective of this approach is twofold: using Panel data gives a clear idea on the exchange rate Pass-Through, and more particularly the long-run one, for some developing countries; on the other hand, non-stationary Panels can point out the existence of a possible cointegration relationship between several variables. This empirical methodology uses the additional information available in the cross-section in order to increase the power of tests to identify non-spurious cointegration between the variables with respect to single country tests; no other study has applied a non-stationary panel cointegration and estimation approach in this context.

The second goal of our paper is to explain the differential impact of change in exchange rate on import prices in some developing countries, in particular in the long run. In this context, several authors analyzed cross-countries difference of exchange rate pass-through into import prices, respectively, Campa and Goldberg (2003) for OECD countries and Webber (1999) for Asian-Pacific countries. Their analyses are based on three macroeconomic determinants: nominal exchange rate volatility, countries openness and inflation. However, developing countries have different characteristics in comparison with the developed countries. This led us to analyse the differences in exchange rate pass-through into import prices across the three following determinants. Firstly, we proxy exchange rate volatility by exchange rate regimes. Secondly, we measure country openness by tariff barriers. Finally, we identify inflation by inflation regimes.

We think that the results of this paper provide more understanding about exchange rate pass-through into import prices in developing countries; this can be used both for international monetary policy and international trade policy.

This paper is organized as follows. In section 2, the price equation is defined. In section 3, the stationarity and cointegration tests are performed. In section 4, by using the appropriate estimation techniques of our long run relation, we show that the long run exchange rate pass-through in developing countries is heterogeneous. In section 5 and 6, in order to explain the cross-country differences in long run exchange rate, we introduce the determinants which can explain the differences in long run exchange rate pass-through in 24 developing countries. We find that countries with fixed exchange rate regimes and lower tariff barriers and higher inflation experience the highest long run exchange rate pass-through into import prices. In section 7, we provide some concluding remarks.

## 2 Exchange rate pass-through equation

Exchange rate pass-through empirical studies were interested in the extent to which exchange rate movements are transmitted to traded goods prices, rather than absorbed in producer profit margins or markups. According to Goldberg and Knetter (1997), exchange rate pass-through is defined as the percentage change in the local currency import prices resulting from a one percent change in the exchange rate between the exporting and importing countries. The exchange rate pass-through into import prices studies are empirically based on a statistical relationship of the elasticity of import prices to exchange rates. Testing this relationship is based on the following equation:

$$\Delta p_t = \gamma \Delta e_t + \varepsilon_t. \tag{1}$$

where $p_t$ and $e_t$ are respectively the natural logarithm of import price and nominal exchange rate; $\varepsilon_t$ is an error term and $\gamma$ is the exchange rate pass-through coefficient. The extent of exchange rate pass-through coefficient is based on the value of $\gamma$. A one to one response of import prices to exchange rate is known as a complete exchange rate pass-through and $\gamma = 1$, while the case where the exchange rate pass-through coefficient is less than 1 ($\gamma < 1$) is known as partial or incomplete exchange rate pass-through.

However, Campa and Goldberg (2003) criticize this specification because it only represents a non-structural statistical relationship and lacks an economic interpretation. They argue that a correct specification should include, additionally, controls to capture exporter's costs associated with local inputs and demand conditions in the destination country. Recent empirical studies[1] on exchange rate pass-through into import prices use an approach based on micro-foundations of pricing behavior by exporting firms.

In this paper, the equation that we use to estimate the degree of the exchange rate pass-through into import prices is similar to the equation used in the literature in this area (Hooper and Mann (1989), Goldberg and Knetter (1997) and Campa and Goldberg (2003)). We consider a representative foreign firm having some degree of control over the price of its goods in an importing country. Assume that this representative firm establishes the price of its exports to country i (i is a developing country) in its own currency ($PX_{it}$) at a markup ($\lambda_{it}$) over its marginal cost of production ($C^*_{it}$), that is:

$$PX_{it} = \lambda_{it} C^*_{it}. \tag{2}$$

The import price in the domestic currency $PM_{it}$ is obtained by multiplying the export price $PX_{it}$ by the exchange rate of the importing country i, $E_{it}$, that is,

$$PM_{it} = E_{it} PX_{it} = E_{it} \lambda_{it} C^*_{it}. \tag{3}$$

[1] Campa and Goldberg (2003) and Eiji (2004).

The markup is assumed to respond to both demand pressure for the exporting country ($Y^*_{it}$) and competitive pressure in the importing country. Competitive pressure in the importing country is measured by the gap between the competitor prices in the importing country market ($P_{it}$) and the production cost of exporting firm. Therefore, according to Hooper and Man (1989) the markup $\lambda_{it}$ is given by

$$\lambda_{it} = \left[\frac{P_{it}}{E_{it}C^*_{it}}\right]^{\alpha} Y^{*\beta}_{it},\ 0 < \alpha < 1,\ \text{and } 0 < \beta < 1. \tag{4}$$

Substituting equation (4) into equation (2), we obtain

$$PM_{it} = (E_{it}C^*_{it})^{1-\alpha}(P_{it})^{\alpha}Y^{*\beta}_{it}. \tag{5}$$

The logarithmic form of the equation (5) is thus

$$pm_{it} = (1-\alpha)e_{it} + (1-\alpha)c^*_{it} + \alpha p_{it} + \beta y^*_{it}, \tag{6}$$

where lowercase letters denote the logarithmic values of the variables.

In equation (6), the exchange rate pass-through, defined as the partial elasticity of import price with respect to exchange rate, is $(1-\alpha)$. One weakness of this equation is that the pass-through of exchange rate and foreign cost into import price are the same. However, in practice, this restriction does not necessarily hold. Indeed, Bache (2002) argues that exchange rates are more variable than costs, and a reasonable conjecture is that exporters will be more willing to absorb into their markups changes in exchange rates than change in costs, which are likely to be permanent. Moreover, Athukorala and Menon (1995) have provided purely economic reasons to justify that the coefficient restrictions may not hold such as the incompatibility of price proxies which may result from differences in aggregation levels and in methods of data collection. Therefore, in the estimation, we relax these restrictions and consider the following equation (the long run relationship ):

$$pm_{it} = \alpha_i + \beta_1 e_{it} + \beta_2 c^*_{it} + \beta_3 p_{it} + \beta_4 y^*_{it} + \varepsilon_{it}. \tag{7}$$

In this equation[2], the marginal cost of production of foreign firm is difficult to measure, therefore we adopt the Wholesale price movements of the major trade partners of country i (see Eiji (2004)) represented by

$$C^*_{it} = Q_{it} \times \frac{\widetilde{P}_{it}}{E_{it}}, \tag{8}$$

where $E_{it}$ is the nominal effective exchange rate of country i, $\widetilde{P}_{it}$ is the wholesale price index and $Q_{it}$ is the real effective exchange rate. Taking the logarithm of each variable, we consider:

$$c^*_{it} = q_{it} - e_{it} + \widetilde{p}_{it}. \tag{9}$$

About the other variables in equation (7), the proxy for domestic competitor's price $P_{it}$ is the Producer Price Index of country i (PPI). As a proxy for the demand pressure $Y^*_{it}$, we use the GDP of country i and, for import price $PM_{it}$, we take the import unit value in domestic currency.

# 3 Data sources and Empirical methodology

## 3.1 Data sources

The main problem in empirical studies on developing countries is data availability. Because of the difficulty to find some variables such as the nominal effective exchange rate, we consider a panel of

[2] $\beta_1$ is the long-run exchange rate pass-through.

24 developing countries. The data are annual and span the period 1980- 2003 (24 years). They are obtained from International Financial Statistics.

## 3.2 Panel unit root tests

As a pre-test for cointegration analysis, we first investigate panel non-stationarity of the variables. Here two types of panel unit root tests are employed, the t-bar test proposed by Im, Pesaran and Shin (2003) (henceforth IPS), and the Hadri test (2000). The former, a panel analogous of Said and Dickey (1984), tests the null hypothesis of non stationarity, while the latter, a panel analogue of Kwiatkowski et al (KPSS, 1992), tests the null hypothesis of stationarity.

The Hadri test has two main advantages when compared with the classical IPS methodology. Firstly, it avoids the lack of power of the unit root-based tests by assuming stationarity under the null hypothesis. Secondly, it is particularly adapted for panel data series with short time dimension, which is the case in this paper. When applying the above two tests, an important problem is the cross section dependence (the error terms between the individual errors can be correlated). To deal with this issue, different approaches have been proposed in the literature. Some authors add time dummies in the regressions. Others like Phillips and Sul (2003) use panel unbiased estimators. One can also remove the "aggregate" effects by substracting cross-section means from the original observations. For our case, we adopt the last alternative and work with demeaned data.[3]

The IPS t-bar is designed to test $H_0$: all individual units have unit roots, against $H_1$: some individual units do not have a unit root. Formally,

$$H_0 : \rho_i = 0 \forall i, \quad H_1 : \exists \text{ i such that } \rho_i < 0,$$

where $\rho_i$ is the coefficient of the Augmented Dickey-Fuller (ADF) regression for each individual unit,

$$y_{it} = \mu_i + \rho_i y_{it-1} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta y_{it-j} + \gamma_i t + \varepsilon_{it}, \text{ t=1,...T}, \tag{10}$$

where $\gamma_i$ could be zero or not. In our case all variables are assigned to $y_t$. Im et al (2003) show that the IPS t-bar consists of the ADF t-values for each individual, and after an appropriate normalization, the IPS t-bar test statistic is asymptotically distributed as N(0,1) under the null hypothesis.

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^{N} t_{\rho_i}. \tag{11}$$

The Hadri test (2000) is based on quite different assumptions on the data generating process of the series. Hadri (2000) proposes a Lagrange multiplier test (LM) based on the residuals; it is a KPSS test applied to panel data based on the following regression:

$$y_{it} = z_{it}^{'} \gamma + r_{it} + \varepsilon_{it}, \tag{12}$$

where $z_{it}^{'}$ is the deterministic component and $r_{it}$ is a random walk:

$$r_{it} = r_{it-1} + u_{it}, \tag{13}$$

where $u_{it} \sim iid(0, \sigma_u^2)$ and $\varepsilon_{it}$ is a stationary process. We can re-write equation (12) as follows

$$y_{it} = z_{it}^{'} \gamma + e_{it}, \tag{14}$$

[3] $\widetilde{y_{it}} = y_{it} - y_{.t}$ where $y_{.t} = \frac{1}{N} \sum_{t=1}^{N} y_{it}$.

where

$$e_{it} = \sum_{j=1}^{t} u_{ij} + \varepsilon_{it}. \tag{15}$$

Note that $\widehat{e_{it}}$ and $\widehat{\sigma_e^2}$ are respectively the residuals and standard error estimations from equation (14) and $S_{it}$ is the residual partial sum

$$S_{it} = \sum_{j=1}^{t} \widehat{e_{ij}}. \tag{16}$$

The LM statistics is:

$$LM = \frac{\frac{1}{N}\sum_{i=1}^{N} \frac{1}{T^2} \sum_{t=1}^{T} S_{it}^2}{\widehat{\sigma_e^2}}. \tag{17}$$

All the test results are shown in Table 1 and 2. We compare the empirical statistics to the critical values given in Table 4 of Im, Pesaran and Shin (2003 ) at the 5% level for N=24 and T=24. We thus conclude that all variables are stationary in first difference. Hadri test results confirm the results found for the IPS tests (all variables are stationary in first difference).

### 3.3 Tests for panel cointegration

Several authors have recently proposed alternative procedures for panel cointegration tests. In order to ensure robustness of results, we employ Pedroni tests. Pedroni (1995, 1999) has developed seven tests based on the residuals from the cointegrating panel regression under the null hypothesis of non-stationarity. They are calculated using the estimated residuals from the following panel regression:

$$y_{it} = \alpha_i \ + \delta_{it} + \gamma_t + X_{it}^{'}\beta_i + e_{it} \ , i = 1, ...., N \ , t = 1, ...., T, \tag{18}$$

where $\beta_i = (\beta_{1i}, \beta_{2i}, ...., \beta_{ni})$ and $X_{it} = (x_{1i,t}, x_{2i,t}, ...., x_{ni,t})$

$$\widehat{e}_{it} = \rho_i \widehat{e}_{it-1} + \xi_{it}. \tag{19}$$

The first four Pedroni tests are based on the within panel estimator, that are known as the Panel Statistics: a variance ratio test (v-statistic), a panel version of the Phillips and Perron (1988) $\rho$-statistic and t-statistic (non-parametric), and the ADF t-statistic (parametric). The null hypothesis is $\rho_i$ =1 against $\rho_i = \rho < 1$. Additional three statistics are based on pooling along the between dimension and they are known as Group Mean Panel Tests. The three Group Mean statistics are extensions of the Phillips and Perron (1998), $\rho$-statistic and t-statistic and a parametric t-statistic. The null hypothesis for this tests is $\rho_i$ =1 against $\rho_i < 1$.The seven tests are asymptotically distributed as normal as follows

$$\frac{K_{NT} - \mu\sqrt{N}}{\sqrt{v}} \sim N(0,1), \tag{20}$$

where $K_{NT}$ is the form of test statistic , $\mu$ and $v$ are respectively the mean and the variance (see table 2 of Pedroni (1999)). As shown in Table 2, all test statistics reject the null of no cointegration.[4]

[4]Except the v-stat, all test statistics have a critical value of -1.64 (if the test statistic is less than -1.64, we reject the null of no cointegration). The v-stat has a critical value of 1.64 (if the test statistic is greater than 1.64, we reject the null of no cointegration).

* We reject the null of no cointegration.

# 4 Long run exchange rate pass-through estimations

## 4.1 PMG and MG Estimation

Previous empirical works that tried to estimate pass-through elasticities, specified equation (7) in first-differences (Campa and Gonzàlez (2002), Campa and Goldberg (2004) and Bailliu and Fujii (2004)). This type of specification allows acquiring estimation of short-run and long-run pass-through. However, in our empirical approach, we need to use a technique that is suitable for dynamic panel data and which allows to take into consideration non-stationarity of variables and cointegration relationship. To better illustrate this point, we use the « Pooled Mean Group estimator » (PMG) proposed by Pesaran, Shin and Smith (2000). The PMG method restricts the long-run coefficients to be equal over the cross-sections, but allows for the short-run coefficients and error variance to differ across groups on the cross-sections. We test for long-run homogeneity using a joint Hausman test[5] based on the null hypothesis of equivalence between the PMG and Mean Group estimator proposed by Pesaran and Smith (1995). The Mean Group estimator is an average of N individual estimations allowing for long-run heterogeneity. If we reject the null, we reject the homogeneity of our cross-section's long-run coefficients.

We estimate the following model:

$$\Delta p_{it}^{im} = \phi_i p_{it-1}^{im} + \beta_i^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta p_{it-j}^{im} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta x_{it-j} + \mu_i + \varepsilon_{it}, \tag{21}$$

where $x_{it}$ is the vector of explanatory variables : $e_{it}, c_{it}^{*}, p_{it}$ and $y_{it}^{*}$ for country i and $\mu_i$ are the fixed effects.

The pooled mean group restriction is that the elements of $\beta$ are common across countries:

$$\Delta p_{it}^{im} = \phi_i p_{it-1}^{im} + \beta^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta p_{it-j}^{im} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta x_{it-j} + \mu_i + \varepsilon_{it}. \tag{22}$$

Estimation could proceed by OLS, imposing and testing the cross-country restrictions on $\beta$. However, this would be inefficient as it ignores the contemporaneous residual covariance. A natural estimator is Zellner's SUR method, which is a form of feasible GLS. However, SUR estimation is only possible if N is smaller than T. Thus Pesaran, Shin and Smith (2000) suggest a maximum likelihood estimator.

In our empirical exploration, we consider two different estimation techniques. First, we restrict all long-run coefficients to be equal over the cross-sections, and second, the homogeneity is imposed only for the long-run pass-through coefficient. In both cases, the Hausman test rejects the assumption of long-run homogeneity.

The PMG and Mean Group estimations for the first case[6] are shown in Table 3. PMG and Mean Group estimates provide a significant short run (0.506)[7] and long-run pass-through coefficients(respectively 0.637 and 0.726). Secondly, by the joint Hausman test, we reject long-run homogeneity with a probability value of 0.03. For the second case, we obtain by PMG estimation a short run coefficient of 0.510 and a long run exchange rate pass-through of 0.789. Mean group estimation

[5] More details will be provided in Appendix.

[6] All PMG and MG estimations were performed using the GAUSS code written by Yongcheol Shin. The program is available on line at http://www.eco.cam.ac.uk/faculty/pesaran/jasa.exe.

[7] Given our data frequency, the short run here refers to one year period.

provide a long run exchange rate pass-through of 0.716. By the Hausman test, we reject the long run homogeneity of exchange rate pass-through coefficients (with a probability value of 0.0056). So, following these results, we conclude that the long run exchange rate pass-through into import prices in developing countries is an heterogeneous phenomenon. Therefore, we are going to use estimation techniques taking into account the heterogeneity of long-run coefficients. We use FMOLS and DOLS between-dimension estimators (Group Mean Estimator) proposed by Pedroni (2001).

## 4.2 Mean Group Panel Estimations

In order to estimate long run coefficients of the cointegration relationship (7), we use FMOLS and DOLS between-dimension estimators (Group Mean Estimator) proposed by Pedroni (2001)[8]. An important advantage of the between-dimension estimators is that the form in which the data are pooled allows for greater flexibility in the presence of heterogeneous cointegrating vectors. Another advantage is that the estimates have a more useful interpretation when the true cointegrating vectors are heterogeneous. Specifically, the estimates for the between dimension estimator can be interpreted as the mean value for the cointegrating vectors, while this is not true for the within-dimension estimates. By analyzing FMOLS and DOLS estimations results, we show that our developing countries experience a higher long-run exchange rate pass-through coefficient. We obtain by FMOLS, an estimation of long-run exchange rate pass-through of 77.2% and by DOLS of 82.7% (see Table 4 ). But the pass-through coefficient is not equal to one [9]. However, the average masks cross-country difference in long run exchange rate pass-through into import prices. For example, by FMOLS, the long-run pass-through coefficients vary from 107% for Algeria ( a complete pass-through coefficient: $\beta_1 > 1$) to 42% for Chile (a partial pass-through coefficient $0 < \beta_1 < 1$) (See Table 5). Similarly, by DOLS the long run pass-through coefficients vary from 110% for Paraguay to 43% for Singapore (See table 6).

There exist several explanations to these cross-country differences of long-run exchange rate pass-through into import prices. Many researchers focussed on theoretical arguments to explain cross-country differences in exchange rate pass-through. Devereux and Engel (2001), and Bacchetta and Van Wincoop (2001) indicate that exchange rate pass-through into import prices can depend on the stability of local monetary policy. If exporters set their prices in the currency of the countries with the stable monetary policies, import prices in local currency terms would be more stable in countries with more stable monetary policies. Country size may be another important factor for heterogeneity of long run exchange rate pass-through into import prices. Dornbusch (1987), stipulates that exchange rate pass-through may be higher if there is a lot of exporters in comparison to the presence of local competitors. Exchange rate pass-through might be inversely related to country real GDP. On the other hand, many empirical analyses[10] have tried to explain cross-country differences in long-run exchange rate pass-through. They focused on many macro determinants such as exchange rate volatility, openness (or country size) of a country, and inflation environment. For our empirical analysis, we focussed on these macro variables in order to explain the differences in long-run exchange rate pass-through into import prices for our 24 developing countries.

[8] Details of these method are available in appendix.

[9] Campa and Goldberg (2003) find that full pass-through is generally supported as a longer run characterization.

[10] More details will be provided in section 5.

# 5 Differences in long run exchange rate pass-through

Once empirically has been tested the long-run exchange pass-through heterogeneity in developing countries, we investigate the potential macroeconomic sources of this long-run heterogeneity. Firstly, we introduce a brief review of some empirical investigations which try to explain the reasons of cross-country differences in exchange rate pass-through into import prices. Secondly, in order to explain the differences in long-run exchange rate pass-through into import prices in developing countries, we use three macroeconomic determinants: exchange rate regimes, tariff barriers and inflation regimes

## 5.1 A brief review of the empirical literature

Several authors have been interested in analyzing the heterogeneity sources of exchange rate pass-through into import prices, such as Campa and Goldberg (2003) for OECD countries, and Webber (1999) for Asian-Pacific countries. Heterogeneity sources used by these authors are nominal exchange rate volatility, the country openness, and inflation.

### 5.1.1 Webber (1999)

In order to estimate short run and long run exchange rate pass-through into import prices for 9 countries of the Asia-Pacific[11], Webber (1999) uses the Johansen (1998) procedure to the following VAR:

$$\Delta Z_t = \sum_{i=1}^{k} \Gamma_i \Delta Z_{t-k} + \Pi Z_{t-1} + \mu + \Psi_1 SD_t + \Psi_2 PED_t + \varepsilon_t, \tag{23}$$

where $Z = [P_{D_t} \ P_{F_t} \ e_t]'$ , $P_D$ is the domestic currency import price, $P_{F_t}$ is the US export price, $e_t$ is the nominal exchange rate, $\mathrm{PED}_t$ refers to a centred pegged exchange rate dummy (used for countries with exchange rates that were pegged against the US$) and $\mathrm{SD}_t$ is a vector of centred deterministic seasonal dummies[12]. To explain cross-country differences in long-run exchange rate pass-through into import prices, Webber (1999) uses the import share of a country to total world import as a country openness indicator, and the average percentage change in absolute terms in the various bilateral exchange rates for seven Asian-pacific countries over the period of observation as a proxy to nominal exchange rate volatility. Webber (1999) expected import shares and exchange rate volatility to be characterized by an inverse relationship with long run pass-through ranking. However, this expectation was not verified for any of the nine countries at the empirical level.

### 5.1.2 Campa and Goldberg (2003)

Campa and Goldberg (2003) used the following log–linear relation :

$$P_t = \alpha + \delta X_t + \gamma E_t + \varphi Z_t + \varepsilon_t, \tag{24}$$

where $\mathrm{P}_t$ is the local currency import prices, $\mathrm{E}_t$ is the exchange rate, $\mathrm{X}_t$ is primary control variable representing exporter costs, and $\mathrm{Z}_t$ is another set of control variables (real GDP for the destination market, among other.).

Then, they estimated the following regression for 25 countries (using quarterly data from 1975 through 1999):

[11] Korea, Thailand, Phillippines, Malaysia, Australia, Japan, Singapore, Pakistan and New Zealand.

[12] For more details, see Webber (1999).

$$\Delta P_t^j = \alpha + \sum_{i=0}^{-4} a_i^j \Delta E_{t-i}^j + \sum_{i=0}^{-4} b_i^j \Delta w_{t-i}^j + c^j \Delta GDP_t^j + \nu_t^j \quad j = 1, ...25. \tag{25}$$

Note that $w_{t-i}^j$ is a proxy for exporter costs to a country j; the estimation methodology is applied to 25 is ordinary least squares. Campa and Goldberg (2003) find an average long run pass-through to import prices equal to 0.77. However, this average masks interesting cross-country differences in exchange rate pass-through into import prices. In order to explain that, Campa and Goldberg (2003) test the significance of some macroeconomic determinants such as country size measured by real GDP[13], exchange rate volatility (the average of the quarterly squared changes in the nominal exchange rate) and average inflation rates by running a regression over the short-run and long-run pass-through elasticities of OECD countries. This regression is given by

$$\gamma^i_{sr \text{ or } lr} = \alpha + \beta x^i + \varepsilon^i, \tag{26}$$

where $x^i$ is a vector representing all the exogenous regressors: country-specific average inflation rates, money growth rates, exchange rate volatility, and real GDP. Campa and Goldberg (2003) find that exchange rate volatility affects in a statistically significant way the degree of pass-through in the short-run: countries with more nominal volatility have higher pass-through into import prices. However, their results show that country size and inflation rate are insignificant in the ranking of long-run exchange pass-through across countries.

## 5.2 Our empirical approach

In order to identify some heterogeneity sources of long-run exchange rate pass-through into import prices in developing countries, we use three macroeconomic determinants used in different empirical approaches such as exchange rate volatility, country openness and inflation. We proxy exchange rate volatility by exchange rate regimes, we measure country openness by trade barriers tariff., finally, we identify inflation by inflation regimes.

### 5.2.1 Exchange rate volatility measured by exchange rate regimes

Developing countries often change their exchange rate regimes. Moreover, since the 90s, several countries opted for floating exchange rate regimes. However, many developing countries, that in theory have a flexible rate, intervene in foreign exchange markets, so that in practice very little difference exists (in terms of observable performance) with countries that have explicit fixed exchange rate regimes. We think that the decomposition into different exchange rate regimes of our panel of countries will be able to give us more information about the nominal exchange rate volatility. There are two approaches to classify countries by exchange rate regimes. The basic reference for the classification of exchange regimes is the International Monetary Fund's Annual Report on Exchange rate Arrangements and Exchange Restrictions (AREAER). This classification is a "de jure" classification based on the publicly stated commitment of the authorities in the country in question. The report captures the notion of a formal commitment to a regime, but fails to capture whether the actual policies were consistent with the stated commitment. The problem of the de jure classification can be solved if the classification is based on the observed behavior of the exchange rate. In this context, Levy-Yeyati and Sturzenegger (2002) propose an alternative classification to the de jure classification: a new de

[13] This measure is meant to capture the extent to which local competitors are large in number relative to foreign firms, which could affect the degree of pass-through, see Dornbusch (1987).

facto classification of exchange rate regimes that reflects actual policies. They define exchange rate regimes according to three classification variables: change in the nominal exchange rate, the volatility of these changes, and the volatility of international reserves. Through these classification variables, fixed exchange rate regimes are characterized by a low volatility of nominal exchange rate and a high volatility of international reserves, while floating exchange rate regimes are associated with high volatility of nominal exchange rate and stable international reserves

- Exchange rate volatility ($\sigma_\varepsilon$), is measured as the average of the absolute monthly percentage changes in the nominal exchange rate during a calendar year.
- The volatility of the exchange rate changes ($\Delta\sigma_\varepsilon$), is computed as the standard deviation of monthly percentage changes in the exchange rate.
- The volatility of international reserves ($\sigma_r$), is measured as the average of the absolute monthly change in international reserves relative to the monetary base in the previous month.

To construct their classification Levy-Yeyati and Sturzenegger use a cluster analysis methodology[14]. This approach allows to classify exchange rate regimes according to some variation boundaries (see table 7).

By using the Levy Yeyati and Sturzenegger classification [15], a classification of our developing countries in two groups is possible (see Table 9).

### 5.2.2 Country Openness measured by tariff barriers

We propose an alternative measure of openness based on the country average tariff rate. This choice is motivated by the fact that most developing countries adopt a trade liberalization that is not accompanied with a significant decrease of trade distortion. By using this criterion (trade distortions that may exist to protect sectors in which import substitution exists), we can have a precise idea about the degree of openness in developing countries. More precisely, our argument is that a country with a low tariff rate is more opened than one with high tariff. Our aim is to classify our 24 countries into different groups according to their degree of openness. Note that it is not obvious to do a classification based on trade tariff. However, the only measures proposed are the TRI (Trade Restrictiveness Index) by Anderson and Neary (1994) or the MTRI (Mercantilist Trade Restrictiveness Index) by Anderson and Neary (2003). The two measures use tariff-barriers and non-tariff-barriers data but we are not able to apply them to our developing countries for the period 1980-2003 because we have not sufficient non-tariff-barriers data for some developing countries. To solve this problem, we use the tariff-barriers data from CEPII (Centre d'Etudes Prospectives et d'Informations Internationales) for the period 1981-2003. Actually, we have the yearly tariff barriers rates for each country of our panel. In order to classify our countries into two groups, we calculate for every year, the median of tariff barriers rate. For each country, we compare the rate applied to this median. All along the period, if a country is in the high group more often than in the low group, then we will consider that this country has higher tariff barriers. Otherwise, we will consider that this country has a lower barriers tariff. This method allows us to classify our developing countries into two groups. We obtain 13 countries with low tariff-barriers and 11 countries with high tariff-barriers (see table 9).

[14] The K-means cluster analysis (KMC)

[15] This classification is available only for period going from 1980 to 2000.

### 5.2.3 Inflation environment measured by inflation regimes

Taylor (2000) argues that the recently-observed declines in the pass-through to aggregate prices are the result of a low inflation environment. A high inflation environment would thus tend to increase exchange rate pass-through. More precisely, the pass-through depends on the policy regime: a credible low inflation regime will automatically achieve a low pass-through.

In order to classify our different countries by regimes of inflation, we use the methodology employed by Choudri and Hakura (2001). The authors analyzed the exchange rate pass-through to domestic prices by taking into account inflationary environment for a sample of 71 countries (including some developing countries) for the period 1973-2000. Choudri and Hakura (2001) classify their countries into three groups: weak inflation, moderate inflation and high inflation. This classification is based on the average of the rate of inflation (mean annual inflation) (see table 8).

According to this classification, Choudri and Hakura (2001) argue that industrial countries are characterized by low inflation while developing countries are characterized by an important variation of the inflation, they are divided in three groups. For our analysis, we divided our 24 developing countries in two regimes: country characterized by mean annual inflation less than 10 % will be considered as low inflation countries, while countries characterized by mean annual inflation higher than 10 % will be considered to be high inflation countries (see table 9).

# 6 Empirical results

We conduct panel unit root tests (Hadri test only)[16] and Pedroni cointegration tests for all groups (exchange rate regimes, trade barriers and inflation regimes). By using the Hadri test (see table 10, 11, and 12), we conclude that most variables are I(1) in level and stationary in first difference. Most of Pedroni's test statistics reject the null of no cointegration.(see table 13) Then, by analyzing FMOLS and DOLS estimation results, we show that

- Countries with fixed exchange rate regimes experience a higher long-run pass-through than floating exchange rate regimes. By FMOLS, we obtain respectively 0.604 and 0.318. DOLS give us respectively 0.626 and 0.379. These results confirm the inverse relationship between long run exchange rate pass-through and nominal exchange rate volatility due to exporters competition for market shares. When exporters face a floating exchange rate regime in the importing country, they will try to maintain their market shares. So, they don't fully pass exchange rate change into import prices (see table 14 and 15).

- Lower tariff barriers countries experience a higher long run exchange rate pass-through than higher tariff barriers . We obtain respectively by FMOLS 0.777 and 0.585. While, we obtain by DOLS 0.834 for lower tariff barriers countries and 0.589 for higher tariff barriers. These results indicate that tariff barriers can explain the cross-country differences of long-run exchange rate pass-through into import prices in developing countries (see table 14. and 15).

- Countries characterized by high inflation regimes experience a higher long run exchange rate pass-through than lower inflation regimes. We obtain respectively by FMOLS 0.425 and 0.395. While, we obtain by DOLS 0.392 for lower inflation regimes and 0.593 for higher inflation regimes. Our results indicate that inflation can explain the cross-country difference of long run exchange rate pass-through difference in developing countries (see table 14 and 15).

[16] This test is more appropriate for lower N and T.

# 7 Concluding remarks

The analysis has focussed on the long-run exchange rate pass-through into import prices on a sample of 24 developing countries by using a non- stationary panel approach. We considered that exchange rate pass-through is determined by a combination of nominal effective exchange rate, the price of competing domestic product, the exporter's cost and domestic demand conditions, leaving aside the hypotheses that developing countries introduce quasi identical economic peculiarities, such us they are hardly dependent on international trade, and that in most cases they are "price takers". We test for long run homogeneity of exchange rate pass-through by using Pooled Mean group approach; Hausman test results lead us to conclude that long exchange rate pass-through is heterogeneous. Then in order to take into account this heterogeneity, we use the Pedroni's Mean group estimators and we find that these countries experience on average a high long run exchange rate pass-through (by FMOLS, we obtain 77.25% and by DOLS, we obtain 82.7%).

Our analysis reveals differences in exchange rate pass-through in our developing countries explained by three macroeconomics determinants: exchange rate regimes, trade distortions and inflation regimes. We show that countries with fixed exchange rate, lower tariff barriers and higher inflation regimes exhibit a higher long-run exchange rate pass-through into import prices than countries with higher tariff barriers, floating exchange rate and lower inflation regimes.

## Acknowledgements

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# Appendix A. Empirical Results

Table 1: **IPS and Hadri panel unit root tests results**

| variables | IPS tests | | | | Hadri tests | |
|---|---|---|---|---|---|---|
| | level | | first difference | | level | first difference |
| variables | intercept | intercept+ trend | intercept | intercept + trend | | |
| pim | -1.252 | -0.932 | -7.420 | -7.896 | 4.928 | 1.329 |
| en | -1.489 | -2.189 | -3.499 | -3.105 | 3.672 | 0.978 |
| ppi | -1.523 | -0.895 | -5.849 | -5.915 | 6.143 | -2.234 |
| $y^*$ | -1.242 | -1.544 | -5.016 | -5.931 | 2.478 | -3.765 |
| c* | -1.439 | -0.955 | -3.792 | -3.162 | 2.326 | 0.765 |

Note: For the IPS tests, the critical value at the 5% level is -1.73 for model with an intercept and -2.45 for model with intercept and linear time trend**.**

Note: For the Hadri tests, the null of stationarity is rejected if the computed Hadri statistic is greater than 1.645 at the 5% level.

Table 2: **Pedroni cointegration tests results**

| *Statistics* | *values* |
|---|---|
| Panel v-stat | $2.438^{**}$ |
| Panel rho-stat | $-2.881^{**}$ |
| Panel pp-stat | $-2.532^{**}$ |
| Panel adf-stat | $-3.238^{**}$ |
| group rho-stat | $-3.409^{**}$ |
| group pp-stat | $-3.166^{**}$ |
| group adf-stat | $-2.304^{**}$ |

Table 3 : PMG and MG estimations

| **1. Homogeneity of all long-run coefficients** | | | | |
|---|---|---|---|---|
| **Estimators** | **PMG** | | **MG** | |
| **variables** | **coefficients** | **t-values** | **coefficients** | **t-values** |
| **en** | **0.637** | **7.968** | **0.726** | **2.434** |
| **ppi** | **0.449** | **11.859** | **0.237** | **2.722** |
| **$y^*$** | **0.454** | **10.841** | **0.393** | **6.021** |
| **c*** | **0.327** | **2.761** | **0.296** | **1.299** |
| **2. Homogeneity of long-run exchange rate pass-through coefficient** | | | | |
| **Estimators** | **PMG** | | **MG** | |
| **variables** | **coefficients** | **t-values** | **coefficients** | **t-values** |
| **en** | **0.799** | **32.471** | **0.716** | **32.71** |
| **ppi** | **0.869** | **31.414** | **0.210** | **3.848** |
| **y*** | **0.768** | **5.868** | **0.530** | **8.584** |
| **c*** | **0.283** | **6.356** | **0.309** | **9.661** |

Table 4: **FMOLS and DOLS Mean Group Panel estimation**

| Estimator | FMOLS | | DOLS | |
|---|---|---|---|---|
| variables | coefficients | t-values | coefficients | t-values |
| en | 0.772 | 2.354 | 0.827 | 6.322 |
| ppi | 0.243 | 5.947 | 0.303 | 4.178 |
| y* | 0.486 | 2.546 | 0.920 | 2.256 |
| c* | 0.286 | 4.178 | 0.291 | 2.234 |

Table 5: **FMOLS estimations by country**

| country | en | ppi | y* | c* |
|---|---|---|---|---|
| 1-Algeria | 1.07 (20.15) | -0.27 (-2.37) | 0.33 (2.27) | -0.40 (-2.77) |
| 2-Burkina Faso | 0.56 (3.25) | -0.35 (-0.42) | -0.68( -1.34 ) | -1.13( -2.26 ) |
| 3-Botswana | 0.37 (4.83) | 0.57 (5.54) | -0.42( -1.86 ) | -1.26( -3.31 ) |
| 4-Ivory Coast | 0.73 (15.63) | -0.28 (-1.15) | 0.02( 0.05 ) | 0.03( 4.18 ) |
| 5-Gabon | 0.43 (2.45) | 0.86 (9.03) | -0.82( -4.11 ) | 0.22( 2.71 ) |
| 6-Morocco | 0.93 (6.20) | 0.73 (1.11) | -0.17( -0.31 ) | 1.65( 1.68 ) |
| 7-Nigeria | 0.64 (5.17) | 0.77 (2.67) | 0.78 (1.10) | -1.20 (-3.06) |
| 8-Senegal | 1.11 (2.71) | -0.10 (-1.69) | 1.14(3.67) | 1.71 (5.44) |
| 9-Tunisia | 0.33 (3.02) | 0.23 (1.10) | 0.02 (10.02) | -0.31 (-0.74) |
| 10-Zambia | 0.88 (10.93) | -0.15 (-11.79) | 1.69 (3.24) | 1.55 (4.85) |
| 11-India | 0.55 (3.03) | 1.34 (5.89) | 3.60 (0.90) | -1.19 (-4.98) |
| 12-Indonesia | 0.29 (2.10) | 0.62 (1.39) | -0.84 (-2.57) | -1.08 (-1.67) |
| 13-Iran | 0.27 (1.55) | 0.51 (5.68) | -1.15 (-2.08) | 0.98 (3.39) |
| 14-Pakistan | 0.47 (4.29) | 0.48 (0.98) | 0.07 (0.45) | -0.13 (-3.32) |
| 15-Phillippines | 0.68 (11.41) | 0.80 (2.01) | 0.18 (0.21) | 0.31 (0.93) |
| 16-Singapore | 0.65 (3.69) | 0.95 (2.42) | 2.32 (2.05) | 2.22 (0.77) |
| 17-Bolivia | 1.17 (3.64) | 1.06 (4.82) | -0.08 (-0.19) | 2.29 (2.41) |
| 18-Chile | 0.42 (6.07) | -0.34 (-5.57) | 0.43 (2.45) | 0.10( 0.76 ) |
| 19-Colombia | 0.74 (4.85) | 4.19 (10.14) | 1.52 (2.46) | 2.73 (2.83) |
| 20-Costa Rica | 2.03 (0.91) | -0.29(-0.19) | 0.68 (0.18) | 0.87 (1.24) |
| 21-Ecuador | 1.21 (1.16) | -1.59(-1.32) | 7.87 (3.22) | 1.38 (3.02) |
| 22-Paraguay | 0.95 (2.69) | -2.10 (-4.98) | -0.18 (-1.06) | 2.72 (6.91) |
| 23-Uruguay | 1.02 (3.98) | 0.38 (2.76) | -1.84 (-6.36) | 0.05( 0.32 ) |
| 24-Venezuela | 1.03 (2.82) | -2.17 (-5.10) | -0.17 (-0.99) | 1.14 (3.27) |

Table 6: **DOLS estimations by country**

| country | en | ppi | $y^*$ | c* |
|---|---|---|---|---|
| 1-Algeria | 1.34 (0.87) | 0.16 (0.29) | 2.52 (8.29) | 0.11 (1.20) |
| 2-Burkina Faso | 0.46 (2.66) | 0.39 (0.53) | 0.32 (1.76) | 4.50 (9.01) |
| 3-Botswana | 0.50 (2.22) | 0.72 (2.11) | 0.42 (2.14) | -0.44 (1.12) |
| 4-Ivory Coast | 1.03 (3.99) | 1.44 (2.89) | 0.05 (0.16 ) | -2.71 (-2.72) |
| 5-Gabon | 0.39 (2.76) | 0.85 (2.43) | 0.12 (0.89) | 0.565 (0.09) |
| 6-Morocco | 1.12 (3.67) | 0.95 (3.19) | 0.13 (0.31) | -0.15 (-4.71) |
| 7-Nigeria | 0.45 (1.68) | 0.61 (0.39) | 0.23 (0.42) | 2.42 (0.20) |
| 8-Senegal | 1.12 (2.54) | 0.76 (1.87) | -0.83 (-3.24) | 0.23 (1.79) |
| 9-Tunisia | 0.39 (3.55) | 0.63 (0.26) | 0.12 (0.58) | 0.33 (1.11) |
| 10-Zambia | 0.42 (2.48) | -0.72 (-0.86) | 1.99 (1.39) | -1.40 (-6.71) |
| 11-India | 0.97 (2.42) | -0.51 (-2.08) | 1.32 (2.01) | -5.64(-0.43) |
| 12-Indonesia | 0.41 (10.38) | 0.58 (0.33) | -0.14 (-0.52) | -0.50 (-3.12) |
| 13-Iran | 0.37 (1.14) | -6.14 (-5.66) | 5.86 (5.42) | 3.59 (1.96) |
| 14-Pakistan | 0.43 (2.37) | 0.94 (3.33) | -0.14 (-2.52) | -0.12 (-0.84) |
| 15-Phillippines | 0.75 (2.11) | 2.88 (1.24) | 3.75 (1.61) | 0.97 (2.13) |
| 16-Singapore | 0.43 (2.08) | 0.68 (2.31) | 0.34 (1.82) | -1.16 (-1.94) |
| 17-Bolivia | 1.63 (1.26) | 0.99 (0.13) | -0.14 (-0.12) | 1.21 (1.06) |
| 18-Chili | 0.42 (2.99) | -0.24 (-0.96) | 1.36 (1.16) | 0.14 (0.18) |
| 19-Colombia | 0.67 (4.70) | 1.43 (0.97) | -0.01 (-0.70) | 0.26 (1.03) |
| 20-Costa Rica | 2.09 (0.44) | -0.60 (-0.30) | 1.73 (0.33) | 0.42 (0.59) |
| 21-Ecuador | 1.03 (2.30) | 1.25 (1.90) | 0.31 (1.09) | 0.80 (1.87) |
| 22-Paragay | 1.10 (3.06) | 0.08 (11.06) | 0.84 (16.05) | 1.67 (0.85) |
| 23-Uruguay | 0.95 (4.25) | 0.07 (2.06) | 0.36 (1.41) | 0.17 (0.19) |
| 24-Venezuela | 1.29 (1.09) | -0.01 (-0.24) | 1.54 (0.07) | 1.09 (0.93) |

Table 7: **Levy-Yeyati and Sturzenegger criterion**

| | criterion 1: $\sigma_\varepsilon$ | | criterion 2: $\Delta\sigma_\varepsilon$ | | criterion 3: $\sigma_r$ | |
|---|---|---|---|---|---|---|
| | min. | max | min | max | min | max |
| **floating** | 0.72% | 2.37% | 0.36% | 1.37% | 0.25% | 6.46% |
| **intermediate** | 0.16% | 1.77% | 0.33% | 1.58% | 5.38% | 10.63% |
| **administrate** | 0.02% | 1.05% | 0.24% | 1.44% | 0.35% | 7.53% |
| **fixed** | 0.00% | 0.63% | 0.00% | 0.66% | 5.65% | 11.02% |

Table 8: **Choudri and Hakura classifications**

| inflation regimes | mean annual inflation |
|---|---|
| low inflation | < than 10% |
| moderate inflation | between 10% and 30% |
| high inflation | > than 30% |

Table 9: **Country classifications by Trade distortions,Exchange rate regimes and inflation regimes**

| **Trade distortions** | | **Exchange rate regimes** | | **Inflation regimes** | |
|---|---|---|---|---|---|
| **low barriers** | **high barriers** | **fixed** | **floating** | **Low inflation** | **High inflation** |
| Indonesia | Algeria | Burkina Faso | Algeria | Algeria | Gabon |
| Singapore | Morocco | Botswana | India | Burkina Faso | Nigeria |
| Ivory coast | Nigeria | Ivory Coast | Nigeria | Botswana | Zambia |
| Gabon | Tunisia | Gabon | Morocco | Ivory Coast | Indonesia |
| Senegal | Zambia | Senegal | Tunisia | Morocco | Iran |
| Colombia | Pakistan | Zambia | Pakistan | Senegal | Bolivia |
| Chile | Burkina Faso | Iran | Indonesia | Tunisia | Chile |
| Uruguay | Iran | Costa Rica | Singapore | India | Colombia |
| Philippines | Venezuela | Ecuador | Chile | Pakistan | Costa Rica |
| Bolivia | India | Paraguay | Colombia | Philippines | Ecuador |
| Costa Rica | Botswana | Venezuela | Uruguay | Singapore | Paraguay |
| Ecuador | | Indonesia | Philippines | | Uruguay |
| Paraguay | | | Bolivia | | Venezuela |
| | | | Zambia | | |

Table 10: **Hadri Panel unit root tests**

| | Fixed exchange rate regimes | | Floating exchange rate regimes | |
|---|---|---|---|---|
| variables | level | first difference | level | first difference |
| en | 1.866 | 0.421 | 1.884 | 0.931 |
| pim | 1.832 | 0.628 | 1.948 | 0.539 |
| ppi | 2.086 | 0.804 | 2.841 | 0.309 |
| y* | 2.066 | 0.474 | 2.352 | 0.895 |
| c* | 0.466 | 0.592 | 3.475 | 0.707 |

Table 11: **Hadri Panel unit root tests**

| | Low tariff | | High tariff | |
|---|---|---|---|---|
| variables | level | first difference | level | first difference |
| **en** | 2.148 | 1.165 | 2.465 | 1.356 |
| **pim** | 1.569 | 0.435 | 1.921 | 0.657 |
| **ppi** | 3.714 | 1.178 | 2.727 | 0.878 |
| y* | 2.645 | 0.976 | 1.801 | 0.536 |
| c* | 2.963 | 0.928 | 2.125 | 0.740 |

Table 12: **Hadri Panel unit root tests**

| | **Low inflation** | | **High inflation** | |
|---|---|---|---|---|
| variables | level | first difference | level | first difference |
| en | 1.837 | 0.354 | 2.725 | 0.818 |
| pim | 1.672 | -0.987 | 2.719 | 0.788 |
| ppi | 0.577 | -3.901 | 0.423 | -2.664 |
| y* | 2.047 | -0.757 | 3.020 | 0.915 |
| c* | 1.911 | 0.967 | 2.980 | 0.306 |

Table 13: Pedroni's cointegration tests

| | **Exchange rate regimes** | | **Trade-barriers** | | **inflation regimes** | |
|---|---|---|---|---|---|---|
| *Statistics* | fixed | floating | Low tariff | Higher tariff | low inflation | high inflation |
| **Panel v-stat** | 0.0017 | 2.4851* | 1.887** | 2.339** | 0.107 | 5.331* |
| **Panel rho-stat** | -2.7142** | -1.0628 | -1.447 | -1.677** | -5.635* | -15.592* |
| **Panel pp-stat** | -2.4792** | -2.1062* | -3.144** | -0.976 | -2.987* | -6.101* |
| **Panel adf-stat** | -2.9005** | -2.8005* | -2.657** | -3.314** | -5.596* | -31.928* |
| **group rho-stat** | -1.9842** | -1.7658* | -2.505** | -2.439** | -0.476 | -30.504* |
| **group pp-stat** | -1.7806** | -2.1509* | -3.480** | -1.153 | -4.876* | -8.186* |
| **group adf-stat** | -0.4661 | -0.9643 | -0.891 | -2.977** | -1.381 | -8.058* |

Table 14: **FMOLS estimations**

| | Exchange Rate Regimes | | Trade barriers | | Inflation Regimes | |
|---|---|---|---|---|---|---|
| | **Fixed** | **Floating** | **Low** | **Higher** | **Low** | **Higher** |
| **variables** | **coef**[17] | **coef** | **coef** | **coef** | **coef** | **coef** |
| **en** | 0.604 (2.944) | 0.318 (2.303) | 0.777 (2.303) | 0.585 (3.338) | 0.395 (8.278) | 0.425 (6.662) |
| **ppi** | 0.931 (17.995) | 1.472 (10.097) | 0.402 (2.225) | 0.767 (3.506) | 0.820 (5.071) | 0.352 (5.695) |
| $\mathbf{y}^*$ | 1.022 (7.063) | 0.889 (6.781) | 0.455 (3.310) | 1.509 (6.304) | 0.503 (3.033) | 0.412 (11.645) |
| **c*** | −0.723 (−12.354) | −1.617 (−9.587) | −0.242 (−1.353) | −0.691 (−2.072) | −0.505 (−5.326) | −0.076 (−2.261) |

Table 15: **DOLS estimations**

| | Exchange Rate Regimes | | Trade barriers | | Inflation Regimes | |
|---|---|---|---|---|---|---|
| | **Fixed** | **Floating** | **Low** | **Higher** | **Low** | **Higher** |
| **variables** | **coef** | **coef** | **coef** | **coef** | **coef** | **coef** |
| **en** | 0.626 (2.037) | 0.379 (2.583) | 0.834 (10.496) | 0.589 (7.609) | 0.392 (3.802) | 0.593 (3.085) |
| **ppi** | 0.932 (3.708) | 0.535 (0.337) | 0.445 (3.212) | 0.867 (33.052 | 1.053 (5.380) | 0.629 (8.403) |
| $\mathbf{y}^*$ | 0.649 (3.708) | 0.565 (0.667) | 0.369 (4.574) | 1.022 (7.063) | 0.831 (4.136) | 0.013 (0.327) |
| **c*** | −0.594 (−7.386) | −1.773 (−8.683) | −0.233 (−8.886) | −0.995 (−4.878) | −0.794 (−6.911) | −0.663 (−2.188) |

[17] In parentheses is reported the t-value.

# Appendix B. Estimation Methods

## B.1. PMG (Pesaran,Shin and Smith (2000))

Pesaran, Shin and Smith (2000) develop a pooled mean group estimator (PMG) for estimating dynamic heterogeneous panel models. The PMG method is an intermediate case between the averaging and pooling methods of estimation. The PMG method restricts the long-run coefficients to be equal over the cross-section, but allows for the short-run coefficients and error variances to differ across group on the cross-section. We obtain pooled long-run coefficients and averaged short run dynamics as an indication of mean reversion. The PMG is based on an autoregressive distributive lag (p, q,. . . ,q) model:

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{it-j} + \sum_{j=0}^{q} \delta_{ij}^{'} x_{it-j} + \mu_i + \varepsilon_{it}. \tag{27}$$

where X$_{it}$ (K×1) is the vector of explanatory variables for group i, $\mu_i$ represents the fixed effects, the coefficients of the lagged dependent variables ($\lambda_{ij}$) are scalars and ($\delta_{ij}$) are (K×1) coefficients vectors. Equation (27) can be re-parameterised as:

$$\Delta y_{it} = \phi_i y_{it-1} + \beta_i^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta x_{it-j} + \mu_i + \varepsilon_{it}. \tag{28}$$

where $\phi_i = -(1 - \sum_{j=1}^{P} \lambda_{ij}), \beta_i = \sum_{j=0}^{q} \delta_{ij}^{'}, \lambda_{ij}^{*} = - \sum_{m=j+1}^{p} \lambda_{ij}$ and $\delta_{ij}^{*} = - \sum_{m=j+1}^{q} \delta_{im}.$

Firstly, we assume that the residuals in equation (28) are iid, with zero mean, variance greater than zero and finite fourth moments. Secondly, the roots of equation (28) must be outside the unit circle. This assumption ensures that $\phi_i < 0$ and hence there exists a long-run relationship between $y_{it}$ and $x_{it}$ defined by

$$y_{it} = -(\frac{\beta_i^{'}}{\phi_i}) x_{it} + \eta_{it}. \tag{29}$$

The long-run homogeneous coefficient is equal to $\theta = \theta_1 = -(\frac{\beta_i^{'}}{\phi_i})$, which is the same across groups. The PMGE uses a maximum likelihood approach to estimate the model and a Newton-Raphson algorithm. The lag length can be determined using the Akaike Information Criteria.

Pesaran, Shin and Smith (2000) propose a Hausman test. This is based on the result that an estimate of the long-run parameters in the model can be derived from the average (mean group) of the country regressions. This is consistent even under heterogeneity. However, if the parameters are in fact homogenous, the PMG estimates are more efficient. Thus we can form the test statistic:

$$H = \widehat{q}^{'} \left[ Var(\widehat{q}) \right]^{-1} \widehat{q} \tilde{\ } \chi_k^2 \tag{30}$$

where $\widehat{q}$ is a (k×1) vector of difference between the mean group and PMG estimates and $Var(\widehat{q})$ is the correponding covariance matrix.

Under the null that the two estimators are consistent but one is efficient, $Var(\widehat{q})$ is calculated as the difference between the covariance matrices of the two underlying parameter vectors. If the poolability assumption is invalid, the PMG estimates are no longer consistent and we fail the test.

## B.2. FMOLS Mean Group Panel Estimator (Pedroni 2001)

Pedroni's estimator is an average-based estimator defined as the average of the conventional panel FMOLS estimator. The estimation involves five steps from the following model:

$$y_{it} = \alpha_i + x_{it}^{'} \beta + u_{it} \text{ i=1,....,N, t=1,....,T.} \tag{31}$$

**Step 1:** The data transformation are:

$y^*_{it} = (\ y_{it} - \overline{y_i})$ and $x^*_{it} = (x_{it} - \overline{x_i})$ where $\overline{y_i} = \frac{1}{N}\sum_{t=1}^{T} y_{it}$ and $\overline{x_i} = \frac{1}{N}\sum_{t=1}^{T} x^{'}_{it}$.

**Step 2:** We estimate equation (31) using the transformed data. Let $\{u_{it}\}$ the estimated residuals and $\varepsilon_{it} = x^*_{it} - x^*_{it-1}$ and denote $\Omega$ and $\Delta$ two estimators of the long-run covariance and the one side long-run covariance martrices of $w_{it} = (u_{it}, \varepsilon^{'}_{it})$.

**Step 3:** We apply the following transformation:

$$y^+_{it} = y^*_{it} - \widehat{\Omega_{u\varepsilon}}\widehat{\Omega_{\varepsilon}}^{-1}\widehat{\Omega_{\varepsilon u}} \quad \text{and} \quad \widehat{\Delta_{\varepsilon u}}^{+} = \widehat{\Delta_{\varepsilon u}} \widehat{- \ \Delta_{\varepsilon}} \ \widehat{\Omega_{\varepsilon}}^{-1}\widehat{\Delta_{u\varepsilon}}.$$

**Step 4**: The FMOLS estimator is given by:

$$\widehat{\beta}_{FMOLS_{\text{i}}} = \left[\sum_{i=1}^{T} x^*_{it}x^{*'}_{it}\right]^{-1}\left[\sum_{i=1}^{T} x^*_{it}y^+_{it} - T\widehat{\Delta_{\varepsilon u}}^{+}\right]. \tag{32}$$

**Step 5**: The Pedroni between FMOLS estimator is the average of the FMOLS estimator (32) computed for each individual, that is:

$$\widehat{\beta}_B = N^{-1}\sum_{i=1}^{N}\widehat{\beta}_{FMOLS_{\text{i}}}. \tag{33}$$

The t-ratio is defined as the average of the t-ratios computed for each individuals of the panel:

$$\widehat{t}_{\widehat{\beta}_B} = N^{-1/2}\sum_{i=1}^{N} t_{\widehat{\beta}_{FMOLS_{\text{i}}}}. \tag{34}$$

## B.3. DOLS Mean Group Panel Estimator (Pedroni 2001)

Pedroni (2001) proposed an estimator based on the average of the panel DOLS estimator "**Group Mean DOLS"** that we can obtain from the following regression:

$$y_{it} = \alpha_i + \beta_i x_{it} + \sum_{k=-K_i}^{K_i}\gamma_{ik}\Delta x_{it-k} + u^*_{it}. \tag{35}$$

We construct the Group- Mean DOLS estimator as :

$$\widehat{\beta}^*_{GD} = \left[N^{-1}\sum_{i=1}^{N}(\sum_{t=1}^{T} Z_{it}Z^{'}_{it})^{-1}(\sum_{t=1}^{T} Z_{it}\widetilde{y}_{it})\right]. \tag{36}$$

where $Z_{it}$ is the 2×(K+1)×1 vector of regressors and $Z_{it} = (x_{it} - \overline{x_i}, \Delta x_{it-k}, ...., \Delta x_{it+k})$.

However, Group Mean DOLS estimator can be constructed by applying the conventional DOLS estimator to the ith member of the panel as follows

$$\widehat{\beta}^*_{GD} = \frac{1}{N}\sum_{i=1}^{N}\widehat{\beta}^*_{Di}. \tag{37}$$

where $\widehat{\beta}^*_{Di}$ is the conventional DOLS estimator applied to the ith member and $\sigma^2_i = \lim_{T\longrightarrow\infty} E\left[T^{-1}(\sum_{t=\&}^{T}\widehat{u}^*_{it})^2\right]$ the long-run variance of the residuals from the DOLS regression (35). The associated t-statistic for the between-dimension estimator can be constructed as

$$\widehat{t}_{\widehat{\beta}^*_{GD}} = N^{-1/2}\sum_{i=1}^{N} t_{\widehat{\beta}^*_{Di}}. \tag{38}$$

where

$$t_{\widehat{\beta}^*_{Di}} = (\widehat{\beta}^*_{Di} - \beta_0)(\widehat{\sigma}_i^{-2} \sum_{t=1}^{T} (x_{it} - \overline{x_i})^2)^{1/2}. \tag{39}$$

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