Long run exchange rate pass-through: Evidence from new panel data techniques

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June 2011

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Abstract

This paper examines the exchange rate pass-through (ERPT) into import prices using recent panel data techniques. For a sample of 27 OECD countries, panel cointegration tests provide an evidence for the existence of long-run equilibrium relationship in pass-through equation. Following Pedroni (2001), we employ both FM-OLS and DOLS estimators and show that long-run ERPT elasticity does not exceed 0.70%. Individual estimates of ERPT are heteregenous across 27 OECD countries, ranging from 0.23% in France to 0.98% in Poland. When we look for the macroeconomic determinants of this long-run heterogeneity, we implement a panel threshold methodology as introduced by Hansen (2000). Our results indicate a regime-dependence of ERPT, that is, countries with higher inflation regime and more exchange rate volatility would experience a higher degree of pass-through.

J.E.L classification: C23, E31, F31, F40

Keywords: Exchange Rate Pass-Through, Import Prices, Panel Cointegration, Panel Threshold

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1 Introduction

The issue of exchange rate pass-through (ERPT) into domestic prices has long been of interest in debates about the conduct of monetary policy and the choice exchange rate regime. By definition, this concept refers to the degree of sensitivity of import prices to a one percent change in exchange rates in the importing nation's currency. It is commonly argued in pass-through literature that the import prices do not move one-to-one following exchange variations, that is, ERPT is found to be incomplete. Moreover, several industrialized countries have experienced decline in pass-through since the early 1990's. However it is still difficult to answer the question of what factors exactly have caused this trend. In fact, there are several explanations for the reducing pass-through mechanism. From a macroeconomic perspective, the moving towards more stable inflation environment has played an important role in the recent fall in ERPT. This positive correlation between inflation and the degree of pass-through has put forth by Taylor (2000). Known as Taylor's hypothesis, this argues that countries with low-inflation environment as a result of more credible monetary policies would experience a reduced degree of pass-through. Thus inflation regime can be considered as one of the sources of ERPT differences across countries. For instance, it is arguable that pass-through is always higher in developing economies with more than one-digit level of inflation.

In fact, there are several factors influencing ERPT that are often discussed in pass-through literature. In addition to the inflation environment, Campa & Goldberg (2005) have tested the importance of other macroeconomic variables that affecting the pass-through, namely, monetary policy stability, country size and exchange rate volatility. The authors found that find that inflation rate and exchange rate volatility affects in a statistically significant way the degree of pass-through. In their study, Choudhri & Hakura (2006) show that ERPT is positively correlated to the average of inflation rate and the inflation and exchange rate volatility, but no significant role for the degree of openness was founded. The present paper follows this strand of literature and, therefore, analyzes the role of some macroeconomic variables that may account for the cross-country differences in pass-through. In a sample of 27 OECD countries, we address the question of whether inflation rate, degree of openness and exchange rate volatility are potential sources of heterogeneity in ERPT. Using panel threshold model, introduced by Hansen (1999), we show that our sample of countries can be classified into different groups according to their macroeconomic *regimes*. This enables us to test the presence of regime-dependence in ERPT mechanism. To the best of our knowledge, this is the first study applying panel threshold method in this context.

Another important issue in the literature concerns the long-run equilibrium in the pass-through equation. In fact, several empirical studies have failed to find evidence of cointegrating relationship in the data. As discussed in panel cointegration literature (Pedroni (1999, 2001, 2004) and Breitung & Pesaran (2005), among others), conventional nonstationary tests have low power in small sample sizes, so adding the cross-section dimension to the time series dimension would increase the power of these tests. Therefore, we propose to use panel data cointegrating techniques to restore the long-run equilibrium in ERPT relationship.

The first goal of our paper is to measure the long-run ERPT into import prices index for 27 OECD countries. We follow Pedroni (2001) methodology by applying FMOLS and DOLS group mean estimators. Little is said about long run pass-through in this context, and the aim of our paper is to fill this gap by using these recent panel data techniques. The second goal is to provide insights into the factors underlying cross-country differences in pass-through elasticities. To this end, we explore three macroeconomic determinants, i.e. inflation rate, degree of openness and exchange rate volatility which are potential sources of heterogeneity in ERPT. To preview our results, we first provide a strong evidence of incomplete ERPT in our panel 27 OECD countries. On the long run, import prices do not move one-to-one following exchange rate depreciation. Both FM-OLS and DOLS estimators show that pass-through elasticity does not exceed 0.70%. When considering individual estimates, we can note a cross-country difference in the long run ERPT. Especially, there is an evidence of complete pass-through for 5 out of 27 OECD countries, namely, Czech Republic, Italy, Korea, Luxembourg and Poland. Second, when split our sample in different country regimes, our results reveal a regime-dependence of ERPT, i.e. countries with higher inflation regime and more exchange rate volatility would experience a higher degree of pass-through. However, we find that ERPT is weakly correlated to the degree of openness.

The remainder of this paper is organized as follows. Section 2 provides an overview of the literature on ERPT and discusses some macro-determinants that may explain cross-country differences in pass-through. Section 3 describes the analytical framework that underlies our empirical specification and the data used in the study. In Section 4, we discuss the empirical methodology used to test stationarity and cointegration in panel. Results of the empirical analysis for our panel of 27 OECD countries as well as for each individual are presented in Section 5. Section 6 discusses some macroeconomic factors determining ERPT. Section 7 concludes.

2 Overview of the literature

Menon (1995) and Goldberg & Knetter (1997) gave a comprehensive review of a large body of empirical literature which deals with the issue of pass-through to import prices. The main finding of this literature is that import prices do not fully respond to a depreciation or appreciation in the domestic currency. This finding is expectable in the short run due to the staggered price setting, and pass-through seems to be much lower than in the longer run. However, price adjustment may be incomplete even in the long run. In a seminal papers, Dornbusch (1987) and Krugman (1987) justifies incomplete pass-through as a result of firms' markup adjustment depending on market destination. Within imperfect competition market, exporters can practice a pricing-to-market (PTM hereafter) strategy by setting different prices for different destination markets[1]. If the firms keep a constant markup, import prices move one-to-one to changes in exchange rates, and there is no evidence of PTM. This latter case refers to denomination of imports in the currency of the exporting country which is called producer-currency pricing (PCP). And if the firm's markup decreases following destination market currency depreciation, PTM occurs and pass-through to import prices is less than complete. When prices do not to vary in the currency of importing country, this refers to local-currency pricing (LCP) strategy and pass-through would be equal to zero.

In a more recent literature, there has been a growing interest in examining the relationship between ERPT and macroeconomic factors. One of the most convincing factors is the inflation environment in each country. Taylor (2000) argued that the responsiveness of prices to exchange rate fluctuations depends positively on inflation. ERPT tends to increase in a higher inflation environment where price shocks are persistent. In this view, a shift towards lower inflation regime, brought about by more credible monetary policies, can give a rise to reduced degree of pass-through. Many empirical studies gave a supportive evidence to this hypothesis, such as Choudhri & Hakura (2006), Gagnon & Ihrig (2004) and Bailliu & Fujii (2004), among others. Another important macroeconomic determinant of pass-through is the exchange rate volatility. In fact, the relative stability of market destination currency plays a substantial role in determining pass-through. Countries with low relative exchange rate variability would have their currencies chosen for transaction invoicing. Thereby, local-currency pricing (LCP) would prevailing and ERPT is less than complete. Thus, we expect that higher import price pass-through would be positively associated with higher exchange rate volatility. Most of pass-through studies find that countries with low nominal exchange rate volatility have a lower

[1] Pricing-to-market is defined as the percent change in prices in the exporter's currency due to a one percent change in the exchange rate. Thus, the greater the degree of pricing-to-market, the lower the extent of exchange rate pass-through.

ERPT. Empirically, Campa & Goldberg (2005) corroborate this positive link in a sample of 23 OECD countries, although microeconomic factors play a much more important role in their study. For the EMU context, Devereux *et al.* (2003) argued that, following the formation of the EMU, the euro would become the currency of invoicing for foreign exporters (LCP). Therefore, European prices will become more insulated from exchange rate volatility and ERPT tend to be lower in such circumstance. Another macro-determinant that have been frequently tested in the empirical literature is the degree of trade openness of a country. One can expect that the more country is open, the higher is price responsiveness to exchange movements. However, results remain mitigate about the relevance of degree of openness. For instance, Choudhri & Hakura (2006) found insignificant role for the import share in their ERPT regression, while McCarthy (2007) provides a weak evidence of a positive relationship between openness and pass-through to import price.

In our empirical, we focus on the ERPT into import prices in the long run, so, from econometric point of view, suitable estimation techniques must be employed. There is a crucial question about the definition of the long measure of pass-through. These are different approaches had been experimented in the empirical literature. One of the most used specifications of the long run ERPT is provided by Campa & Goldberg (2002, 2005). In these studies, the long run elasticity of pass-through is given by the sum of the coefficients on the contemporaneous exchange rate and four lags of exchange rate terms. According to de Bandt *et al.* (2007), this measure is, in some extent, arbitrary and more accuracy long run pass-through must be defined. By using nonstationary panel data techniques, their study propose to restore the cointegrated long run equilibrium in pass-through relationship (see Table 1). As we mentioned above, there has been an increasing use of unit root and cointegration analysis in the context of panel data. This is not surprising as panel techniques can overcome the size and power constraints associated with the use of a single time series[2].

[2]It's well-known that unit root tests have low power in small sample sizes, so adding the cross-section dimension to the time series dimension increase the power of these tests.

Table 1: Main ERPT studies using panel cointegration approach

STUDY	DATA	METHOD	FINDINGS
Barhoumi (2006)	Annual data (1980-2003) for 24 developing countries	Measuring long run ERPT to import prices using panel data cointegration techniques. FMOLS and DOLS between-dimension estimators (Pedroni (2001)).	A higher group mean long-run ERPT coefficient: 77.2% by FMOLS, and 82.7% by DOLS. Cross-country difference in long run ERPT: by FMOLS, coefficients vary from 107% for Algeria to 42% for Chile, and by DOLS, ERPT vary from 110% for Paraguay to 43% for Singapore. Differences in ERPT are due to three macroeconomics determinants: exchange rate regimes, trade barriers and inflation regimes.
Holmes (2006)	Monthly data (1972:4- 2004:6) for 12 European Union countries	Estimation of long-run ERPT to consumer prices using DOLS between-dimension approach.	The ERPT to European Union consumer prices has declined. This decline has occurred against a background of several factors that enhanced the credibility of a low inflation regime.
de Bandt *et al.* (2007)	Disaggregated monthly data (1995-2005) for 1-digit SITC sectors for 11 euro area countries	Different panel data techniques to test for cointegration in the ERPT equation: -First generation panel cointegration tests with no cross-unit interdependence and no breaks (Pedroni (1999)) -Second generation tests with a factor structure for cross-section dependence and allowing for an individual structural break (Banerjee & Carrion-i Silvestre (2006)).	Commodity sectors (SITC 2 and SITC 3) tend to have a higher (closer to 1) pass-through than manufacturing sectors. Strong evidence of a change in the long run ERPT behavior around the formation of the Economic and Monetary Union (EMU) or close to the period of appreciation of the euro in 2001. Long run ERPT has generally increased after these break dates especially for Italy, Portugal and Spain.
Holmes (2008)	Annual data (1971-2003) for 19 African countries.	FMOLS procedure is employed to obtain long run ERPT to import prices. Using moving window approach to test changing ERPT over time.	Long run ERPT elasticity is about 60% for the African economies. According to moving window estimates, African import prices becoming less sensitive to movements in the exchange rate over time. Decline in the long-run pass-through is accompanied by decreasing in inflation rates occurring since the mid-1990s.

One of the most important economic theories usually tested in this context is the purchasing power parity, for which it is natural to think about long-run properties of data. However, there is a few numbers of studies has investigated the ERPT relationship within a panel data cointegration framework. In Table 1, we summarized the main findings of major studies in this area, namely Barhoumi (2006), de Bandt *et al.* (2007) and Holmes (2006, 2008). Regarding to country's sample, our study is close to those of de Bandt *et al.* (2007) and Holmes (2006), which deal with some countries of the European Union. Nevertheless, our sample is larger since we consider 27 OECD countries in our empirical work. Also, our country sample is more heterogeneous than the listed studies, so using Pedroni (2001) approach is relevant since it allows the long-run cointegration relationships to be heterogeneous across countries.

3 Analytical framework and Data description

3.1 Pass-Through Equation

Our approach is to use the standard specification used in the pass-through literature as a starting point (Goldberg & Knetter (1997) and Campa & Goldberg (2005)). By definition, the import prices, MP_{it}, for any country i are a transformation of the export prices, XP_{it}, of that country's trading partners, using the nominal exchange rate, E_{it} (domestic currency per unit foreign currency):

$$MP_{it} = E_{it}.XP_{it} \tag{1}$$

Using lowercase letters to reflect logarithms, we rewrite equation (1):

$$mp_{it} = e_{it} + xp_{it} \tag{2}$$

Where the export price consists of the exporters marginal cost, MC_{it} and a markup, $MKUP_{it}$:

$$XP_{it} = MC_{it}.MKUP_{it} \tag{3}$$

In logarithms we have:

$$xp_{it} = mc_{it} + mkup_{it} \tag{4}$$

So we can rewrite equation (2) as:

$$mp_{it} = e_{it} + mc_{it} + mkup_{it} \tag{5}$$

Markup is assumed to have two components: (i) a specific industry component and (ii) a reaction to exchange rate movements:

$$mkup_{it} = \alpha_i + \Phi e_{it} \tag{6}$$

Exporter marginal costs are a function of the destination market demand conditions, y_{it}, and wages in exporting country, w^*_{it}:

$$mc_{it} = \eta_0 y_{it} + \eta_1 w^*_{it} \tag{7}$$

Substituting (6) and (7) into (5), we derive:

$$mp_{it} = \alpha_i + \underbrace{(1+\Phi)}_{\beta} e_{it} + \eta_0 y_{it} + \eta_1 w^*_{it}, \tag{8}$$

The structure assumes unity translation of exchange rate movements. This empirical setup permits the exchange rate pass-through, represented by $\beta = (1+\Phi)$, to depend on the structure of competition in one industry. Exporters of a given product can decide to absorb some of the exchange rate variations instead of passing them through to the price in the importing country currency. So if $\Phi = 0$, the pass-through is complete and their markups will not respond to fluctuations of the exchange rates. This is the case when import prices are determined in the exporter's currency (PCP is prevailing). And if $\Phi = -1$, exporters decide not to vary the prices in the destination country currency and, thus, they fully absorb the fluctuations in exchange rates in their own markups (LCP is prevailing).

Thus the final equation can be re-written as follows:

$$mp_{it} = \alpha_i + \beta e_{it} + \gamma y_{it} + \delta w_{it}^{*} + \varepsilon_{it}, \tag{9}$$

The most prevalent result is an intermediate case where ERPT is incomplete (but different from zero), resulting from a combination of LCP and PCP in the economy. So, there is a fraction of import prices are set in domestic currency, while the remaining prices are set in foreign currency. Thus, the extent to which exchange rate movements are passed-through prices will depend on the predominance of LCP or PCP: the higher the LCP, the lower the ERPT, and the higher PCP, the higher ERPT.

3.2 Data description

In this study, we consider the following panel of 27 OECD countries: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Korea, Luxembourg, Netherlands, Norway, New Zealand, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom and United States. The data are quarterly and span the period 1994:1-2010:4. We use price of non-commodity imports of goods and services imports from OECD's Main Economic Outlook as a measure of the import prices, mp_{it}. From the same Data base we take the real GDP as proxy for the domestic demand, y_{it}. To capture movements in the costs of foreign producers, W_{it}^{*}, that export to the domestic market, we use the same proxy adopted by Bailliu & Fujii (2004) represented by: $W_{it}^{*} = Q_{it} \times W_{it}/E_{it}$, where E_{it} is the nominal effective exchange rate (domestic currency per unit of foreign currencies)[3], W_{it} is the domestic unit labor cost and Q_{it} is the real effective exchange rate. Due to data availability, we follow Campa & Goldberg (2005) by using consumer price index, P_{it}, to capture movement in production costs, assuming that prices move one-to-one to shift in wages. Taking the logarithm of each variable, we obtain the following expression: $w_{it}^{*} = q_{it} - e_{it} + p_{it}$. Since nominal and real effective exchange rate series are trade weighted, this gives us a measure of trading-partner costs (over all partners of importing country), with each partner weighted by its importance in the importing country's trade. Data used to construct foreign producers costs - nominal effective exchange rate, consumer prices index and real effective exchange rate - are obtained from IMF's International Financial Statistics.

[3] Home-currency depreciations appear as increases in the nominal effective exchange rate series.

4 Unit root and panel cointegration tests

Before testing for a cointegrating relationship, we investigate panel non-stationarity of the variables included in equation (9). We use the t-bar test proposed by Im *et al.* (2003) (henceforth IPS), which tests the null hypothesis of non stationarity. This test allows for residual serial correlation and heterogeneity of the dynamics and error variances across groups. The t-bar statistic constructed as a mean of individual ADF statistics and is designed to test the null that all individual units have unit roots:

$$H_0 : \rho_i = 0,\ \forall i$$

Against the alternative that at least one of the individual series is stationary:

$$H_1 : \begin{cases} \rho_i < 0 & \text{for } i = 1,2,...,N_1 \\ \rho_i = 0 & \text{for } i = N_1, N_2, ..., N \end{cases} \quad \text{with } 0 < N_1 \leq N$$

Where ρ_i is the coefficient of the Augmented Dickey-Fuller (ADF) regression for each individual unit,

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

As we mentioned above, the IPS t-bar statistic is defined as the average of the individual ADF statistic, $t_{\rho i}$, and tends to a standard normal distribution as $N, T \to \infty$ under the null hypothesis:

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

IPS tests results are shown in Table 2, for both levels and first differences and with different deterministic components. In the level case, we are unable to reject the null hypothesis that all series are non-stationary in favor of the alternative hypothesis that at least one series from the panel is stationary. For tests on the first differences, we can see that the non-stationary null is rejected at the 5% significance level or better. We thus conclude that all variables are stationary in first difference[4].

[4]We compare the empirical statistics to the critical values given in Table 2 of Im *et al.* (2003) at the 5% level for $N = 25$ and $T = 70$.

Table 2: Results for Im, Pesaran and Shin's (2003)

Variables	**Level**		**First difference**	
	Intercept	Intercept & trend	Intercept	Intercept & trend
mp_{it}	-0.5301	-1.3952	-14.0574	-17.2283
e_{it}	0.6973	-0.1440	-10.3162	-11.1541
y_{it}	-1.1128	1.0780	-7.5770	-11.0068
w^{*}_{it}	-1.1586	-0.9063	-10.3704	-15.7305

Note: For the IPS tests, the critical value at the 5% level is -1.81 for model with an intercept and -2.44 for model with intercept and linear time trend. Individual lag lengths are based on Akaike Information Criteria (AIC).

In order to check the long run cointegrating pass-through relation, we employ Pedroni (1999) residual-based tests. Like the IPS panel unit-root test, Pedroni's methodology take heterogeneity into account using specific parameters which are allowed to vary across individual members of the sample. Pedroni (1999) has developed seven tests based on the residuals from the cointegrating panel regression under the null hypothesis of non-stationarity. The first four tests (*panel v-stat, panel rho-stat, panel pp-stat, panel adf-stat*) are based on pooling the data along the within-dimension that are known as the *panel cointegration statistics*. The next three tests (*group rho-stat, group pp-stat, group adf-stat*) are based on pooling along the between-dimension and they are denoted *group mean cointegration statistics*. All tests are calculated using the estimated residuals from the following panel regression:

$$\begin{aligned} y_{it} &= \alpha_i + \delta_{it} + \beta_{1i}x_{1it} + \beta_{2i}x_{2it} + \ldots + \beta_{Ki}x_{Kit} + \varepsilon_{it}, \\ &\quad i = 1,\ldots,N,\ t = 1,\ldots T,\ k = 1,\ldots,K \end{aligned} \tag{12}$$

In fact, both sets of test verify the null hypothesis of no cointegration:

$H_0 : \rho_i = 1,\ \forall i$

Where, ρ_i is the autoregressive coefficient of estimated residuals under the alternative hypothesis ($\hat{\varepsilon}_{it} = \rho_i \hat{\varepsilon}_{it-1} + u_{it}$). We should note that the alternative hypothesis specification is different between the two sets of test:

- The *panel cointegration statistics* impose a common coefficient under the alternative hypothesis which results:

$H_1^w : \rho_i = \rho < 1,\ \forall i$

- The *group mean cointegration statistics* allow for heterogeneous coefficients under the alternative hypothesis and it results:

$$H_1^b : \rho_i < 1, \forall i$$

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

$$\frac{\chi_{NT} - \mu\sqrt{N}}{\sqrt{\upsilon}} \rightarrow N(0,1), \tag{13}$$

Where, χ_{NT}, is the statistic under consideration among the seven proposed, μ, and, υ, are respectively the mean and the variance tabulated in Table 2 of Pedroni (1999). As shown in Table 3, all test statistics reject the null of no cointegration.

Table 3: Pedroni (1999) Cointegration Tests Results

Tests	**1994:1 - 2010:4**
Panel v-stat	6.93854**
Panel rho-stat	-6.20244**
Panel pp-stat	-6.60297**
Panel adf-stat	-5.01230**
Group rho-stat	-5.18729**
Group pp-stat	-6.63478**
Group adf-stat	-4.72966**

Note: 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).

5 Long run ERPT estimates

Following Pedroni (2001), we employ estimation techniques taking into account the heterogeneity of long-run coefficients. Therefore, FMOLS and DOLS Group Mean Estimator can be used to obtain panel data estimates for long run ERPT. These estimators correct the standard pooled OLS for serial correlation and endogeneity of regressors that are normally present in a long-run relationship. In our empirical analysis, we emphasis on between-dimension panel estimators. It's worth noting that the between-dimension approach allows for greater flexibility in the presence of heterogeneity across the cointegrating vectors where pass-through coefficient is allowed to vary[5]. Additionally, the point estimates of the between-dimension estimator can be interpreted as the mean value of the cointegrating vectors, while this is not the case for the within-dimension estimates[6]. To check robustness of our result, we also reporting estimation results for fixed-effects estimators.

According to Table 4, long run pass-through coefficient is statistically significant with the expected positive sign, and the results are fairly robust across estimation techniques. For instance, FM-OLS estimator suggests that one percent depreciation of the nominal exchange rate increases import prices by 0.67%. As we mentioned above, pass-through equation (9) assume unity elasticity of import prices to exchange rate movements in order to account for complete ERPT. However, the null of unity pass-through coefficient ($H_0 : \beta = 1$) is strongly rejected through the different econometric specifications (see t-statistics reported between square brackets in Table 4).

Table 4: Panel Estimates For 27 OECD countries over 1994:1-2010:4

	Dependent Variable: Import Price Index		
	Group mean FM-OLS	**Group mean DOLS**	**Fixed effects**
e_{it}	0.67***	0.69***	0.70***
	(30.21)	(26.69)	(33.01)
	[16.71]	[16.89]	[10.29]
y_{it}	0.27***	0.20***	0.23***
	(6.15)	(6.40)	(11.86)
w^*_{it}	0.68***	0.71***	0.214***
	(7.09)	(6.89)	(8.215)

Note: Group mean FM-OLS and DOLS estimators refer to between-dimension. These estimates include common time dummies. *** indicate statistical significance at the 1 percent level. Pass-through estimates are accompanied by two t-statistics. The t-statistics in parentheses are based on the null of a zero ERPT coefficient (H_0: $\beta = 0$). The t-statistics in square brackets are based on the null of unitary elasticity (H_0: $\beta = 1$).

[5]Under the within-dimension approach pass-through elasticity would be constrained to be the same value for each country under the alternative hypothesis.

[6]According to Pedroni (2001), the between-group FMOLS and DOLS estimators has a much smaller size distortion than the within-group estimators.

This is an evidence of incomplete ERPT in our sample of 27 OECD countries. On the long run, import prices do not move one-to-one following exchange rate depreciation. These results are in line with estimates in the literature of exchange rate pass-through into import prices for industrialized countries. For 23 OECD countries, Campa & Goldberg (2005) find that the average of long run ERPT is 0.64%. In this study, producer-currency pricing (or full pass-through) assumption is rejected for many countries. Using panel cointegration analysis, Barhoumi (2006) and Holmes (2008) reject the pass-through unity for developing countries. In accordance with the conventional wisdom that ERPT is always higher in developing than in developed countries, thus, a partial import prices it is expectable for OECD countries. One can think that pass-through would be complete in the long run due to the gradual full adjustment of prices (as sticky prices tend to be a short run phenomenon)[7]. Nevertheless, the pricing behavior of foreign firms can prevent import prices variations following an exchange rate change. Exporters of a given product can decide to absorb some of the exchange rate variations instead of passing them through to the price in the importing country currency. Empirically exchange rates are found to be much more volatile than prices, and then pass-through would be incomplete even in the long run. This finding is in line with the theoretical price discrimination models which assume a degree of pass-through lower than one even in the long run, as a result of PTM.

When considering individual estimates for our 27 countries, we can note a cross-country difference in the long run ERPT masked by the panel mean value. According to Table 5, FM-OLS estimates indicate the highest import prices reaction in Poland with 0.98% followed by Czech Republic with 0.95%. The lowest degree of pass-through is recorded in Denmark and France with the same elasticity of 0.28%. We can note that results are not significantly different from zero for a few numbers of countries, but it is important to mention that there is an evidence of complete pass-through for 5 out of 27 countries, namely Czech Republic, Italy, Korea, Luxembourg and Poland. This is partly corroborating Campa & Goldberg (2005) results for which producer-currency pricing (PCP) are accepted for Poland and Czech Republic. Moreover, we can observe a low ERPT in the United States by 0.38%, which is a common result in the literature. For example, Campa & Goldberg (2005) find 41% US pass-through elasticity.

[7] For example, see Smets and Wouters (2002).

Table 5: Long run individual ERPT for 27 OECD Countries

Results from FM-OLS method			
Country	**FM-OLS**	t-stat for H_0: $\beta = 0$	t-stat for H_0: $\beta = 1$
Australia	0,78*	32,7	9,04
Austria	-0,08	-0,23	3,28
Belguim	-0,04	-0,28	6,57
Canda	0,76*	18,42	5,75
Switzerland	0,39*	3,32	5,14
Czech Republic	$0,95^{*\#}$	10,75	0,54
Germany	0,63*	4,2	2,44
Denmark	0,28*	3,82	4,05
Spain	0,62*	4,16	2,54
Finland	-0,19	-1,49	9,53
France	0,28*	2,13	5,41
United Kingdom	0,45*	7,24	8,71
Greece	-0,11	-0,45	4,69
Ireland	0,14	1,45	8,7
Iceland	0,66*	11,44	6
Italy	$0,73^{*\#}$	5,25	1,92
Japan	0,44*	4,15	5,28
Korea	$0,87^{*\#}$	7,34	1,12
Luxembourg	$0,85^{*\#}$	2,44	0,43
Netherlands	0,17	1,87	9,17
Norway	0,53*	5,02	4,43
New Zealand	0,85*	16,83	2,98
Poland	$0,98^{*\#}$	8,01	0,14
Portugal	-0,1	-0,27	2,97
Slovak Republic	0,07	0,39	5,13
Sweden	0,48*	5,77	6,23
United States	0,38*	9,71	16,08
Mean Group panel estimation	0,67*	30,21	16,71

Note: *(#) implies that ERPT elasticity is significantly different from 0 (1) at the 5% level. Column (2) reports t-stat for H_0: $\beta = 0$ and column (3) reports t-stat for H_0: $\beta = 1$.

6 Macroeconomic Factors Affecting Pass-Through

Cross-country differences in import prices adjustment would raise the question of what are the underlying determinants of pass-through. To provide an explanation for this long run heterogeneity, we examine some macroeconomic factors that may affect pass-through. Three main factors are selected for this purpose: inflation rates measured as the year-on-year quarterly inflation rate; degree of openness as the percentage of import share in domestic demand; and volatility of exchange rate changes, $\sigma_{\Delta e}$, proxied by the standard deviation of quarterly percentage changes in the nominal effective exchange rate. A summary of the average of these macroeconomic variables over 1994-2010 is given in Table 6. The aim of our analysis is to link those factors to the extent of pass-through. To achieve this, we try to split our panel of countries into different groups with respect to each macroeconomic criteria, and then estimate the ERPT for those different groups. The idea is to compare pass-through elasticity for different country regimes and to draw conclusion about the reasons of cross-country differences in ERPT into import prices.

Table 6: OECD Countries Statistics (1994-2010)

Country	Inflation Rate (%)	Openness (%)	Exchange Rate Volatility (%)
Australia	2,7	16,9	8,2
Austria	1,8	43,6	8,1
Belguim	1,8	66,1	8,6
Canda	2	38,8	4,6
Switzerland	0,9	40,2	7,6
Czech Republic	4,6	75,0	9,4
Germany	1,5	31,3	8,6
Denmark	2,1	38,6	8,2
Spain	3,1	26,4	10,7
Finland	1,4	34,5	13
France	1,6	24,1	8,2
United Kingdom	1,7	27,9	7,1
Greece	4,3	33,5	7,4
Ireland	3,7	67,4	8,3
Iceland	3,2	34,8	14,5
Italy	2,6	23,7	10,5
Japan	-0,1	9,7	8,2
Korea	3,5	32,5	13,2
Luxembourg	2	120,1	9
Netherlands	2,1	55,0	8,7
Norway	2,2	25,7	7,5
New Zealand	2	31,7	10,2
Poland	8,4	33,0	14,7
Portugal	3	35,0	9,2
Slovak Republic	6,7	77,2	9,9
Sweden	1,2	37,4	10,9
United States	2,6	13,9	5,2
Average	**2,7**	**40,5**	**9,2**

Note: Inflation rates measured as the mean of year-on-year quarterly inflation rate over 1994-2010; degree of openness as the mean of import share (% of domestic demand) over 1994-2010; and volatility of the exchange rate changes, $\sigma^{\Delta e}$, is computed as the standard deviation of quarterly changes in the nominal effective exchange rate (average over ver 1994-2010.

Our methodology is close to Choudhri & Hakura (2006) and Barhoumi (2006) studies. Choudhri & Hakura (2006) classify their 71 countries into three groups based on the average of inflation rate. In their study, low, moderate and high inflation groups are defined as consisting of countries with average inflation rates less than 10%, between 10 and 30% and more than 30%, respectively. Similarly, Barhoumi (2006) divided a sample of 24 developing countries between high and low inflation regimes, depending on whether inflation rate is smaller or larger than 10%. However, country classification in these studies is somewhat arbitrary, in the sense that the authors used an ad hoc method to select their sample splits. In our paper, we propose to use panel threshold techniques, introduced by Hansen (1999), to deal with the sample split problem. This methodology enables us to divide our

27 OECD countries into classes based on the value of each "macro-variables", i.e. inflation rate, degree of openness and exchange rate volatility. To the best of our knowledge, the present paper is the only study that applying panel threshold method in this context.

6.1 A single panel threshold model

Hansen (1999) introduce a panel threshold model for a single and multiple threshold levels. Due to our small number of cross sections (27 countries), we consider the single threshold model, so that the observations can be split into two regimes depending on whether the threshold variable is above or below some threshold value. Following Hansen (1999), we can rewrite our pass-through equation as follow:

$$mp_{it} = \begin{cases} \alpha_i + \beta_1' x_{it} + \varepsilon_{it}, & q_{it} \leq \theta, \\ \alpha_i + \beta_2' x_{it} + \varepsilon_{it}, & q_{it} > \theta. \end{cases} \tag{14}$$

Another representation of (14) which is often used in threshold model literature is:

$$mp_{it} = \alpha_i + \beta_1 x_{it} I(q_{it} \leq \theta) + \beta_2 x_{it} I(q_{it} > \theta) + \varepsilon_{it} \tag{15}$$

The dependent variable of our ERPT panel threshold model is the import prices, mp_{it}, and the explanatory variables - Exchange rate, domestic demand and foreign costs - are included the vector $x_{it} = (e_{it}, y_{it}, w^*_{it})'$. $I(.)$ is an indicator function, α_i denotes the level of country i fixed-effect and ε_{it} is a zero mean, finite variance, i.i.d. disturbance. The two regimes are distinguished by different regression slopes, β_1 and β_2, depending on whether the threshold variable q_{it} is smaller or larger than a threshold θ. If the threshold variable q_{it} is below or above a certain value, θ, then the vector of exogenous variable x_{it} has a different impact on the dependent variable, mp_{it}, with $\beta_1 \neq \beta_2$. The threshold variable q_{it} may be an element of x_{it} or a variable external to model. Effectively, in our implementation of the threshold panel method, we consider three different threshold variables - inflation rate, π_{it}, degree of openness, $open_{it}$, and exchange rate volatility, $\sigma^{\Delta e}_{it}$ - which are not belonging to explanatory variables of the pass-through equation. Thus, we will estimate equation (28) for our different threshold variables, $q_{it} = \pi_{it}, open_{it}, \sigma^{\Delta e}_{it}$.

The determination of the estimated threshold, $\hat{\theta}$, is based on two steps procedure using ordinary least squares (OLS) method[8]. In the first step, for any given

[8]Estimation techniques for panel threshold model is given in the appendix A.3 with more details.

threshold, θ, the sum of square errors is computed separately. In the second step, by minimizing of the sum of squares of errors, $S_1(\theta)$, the estimated threshold value, $\hat{\theta}$ is obtained and the residual variance, $\hat{\sigma}^2$, is saved. To check whether the threshold is in fact statistically significant, the null hypothesis of no threshold effect is tested: $H_0 : \beta_1 = \beta_2$. The likelihood ratio test of H_0 is based on the following F-statistics: $F_1 = (S_0 - S_1(\hat{\theta}))/\hat{\sigma}^2$, where S_0 and $S_1(\hat{\theta})$ are sum of squared errors under null and alternative hypotheses, respectively. The asymptotic distribution of F_1 is non-standard. Hansen (2000) propose to use a bootstrap procedure to compute the p-value for F_1 under H_0. Once a significant single threshold is found, we can estimate the pass-through coefficient for each regime. For the purpose of our analysis, we use the estimated threshold to divide our country sample into different groups with respect to their macroeconomic environment (inflation level, degree of openness and exchange rate volatility)[9]. Then, we estimate the ERPT elasticity for each class of countries in order to make a comparison between different regimes.

6.2 Estimation of a single threshold

The estimation results of the threshold levels for each of our macro-determinant are reported in Table 7. Also, we give the plots of sum of squared residuals for the different threshold variables (see Appendix C). When we consider inflation rate as threshold variables, $(q_{it} = \pi_{it})$, we find a threshold level close to 2% $(\theta_\pi = 0,019)$. The test for a single threshold is significant with a bootstrap p-value of 0,04. Given this threshold value, we can define two groups of countries based on *inflation-regime*, i.e. with respect to the average of inflation rate. Thus, we consider countries with mean of inflation equal or less than 2% as *low inflation* countries, while countries with inflation mean more than 2% as *moderate inflation* countries[10]. According to this classification, we obtain 12 low inflation countries and 15 countries with moderate inflation-regime (see Table 8).

The next threshold variable considered in pass-through equation is the degree of openness. According to Table 7, the estimated threshold value is 31,8% of import share, but the presence of a single threshold is non significant according to bootstrapped p-value (0,26). Nevertheless, this threshold value still the best point to consider to split our sample with the respect to the degree of openness (see figure 2 in Appendix C). Thus, we will consider countries characterized by degree of openness less than 32% as *less open* countries, while countries having import share

[9]We follow the same strategy of Hansen (2000) who used the threshold values to split his sample of 565 US firms into *low debt* and *high debt* firms.

[10]The term of *moderate inflation* is used instead of *high inflation* since we don't have double-digit inflation countries in our sample of 27 OECD countries.

larger than 32% will be defined as *more open* countries. This gives us 10 less open countries and 17 countries with degree of openness more than 32%.

Finally, the last criterion which can explain differences in pass-through elasticity is the exchange rate volatility. Different sort of proxies are used in the ERPT literature. For instance, Campa & Goldberg (2005) take the average of the monthly squared changes in the nominal exchange rate. For McCarthy (2007) exchange rate volatility is measured by the variance of the residuals from the exchange rate equation in the VAR. In our empirical analysis, we adopt the same exchange rate volatility proxy employed by Barhoumi (2006) and compute exchange rate volatility as the standard deviation of quarterly percentage changes in the exchange rate, $\sigma^{\Delta e}$ [11]. According to Hansen's single threshold test, we find a significant threshold value equal to 0,082 (see Table 7). Accordingly, we will call countries for whom the mean of exchange rate volatility is less than 8.2% as *less volatility* countries, and the sub-sample of countries having $\sigma^{\Delta e}$ more than 8.2% as *high volatility* countries. We count 11 low volatility countries and 16 high volatility countries (see Table 8).

Table 7: Hansen (1999) test for a single threshold

$$mp_{it} = \alpha_i + \beta_1 x_{it} I(q_{it} \leq \theta) + \beta_2 x_{it} I(q_{it} > \theta) + \varepsilon_{it}$$

Inflation rate: $(q_{it} = \pi_{it})$	
Threshold value $(\hat{\theta}_\pi)$	0.019
F-test	168.03
Bootstrapped p-values	(0.040)
Degree of openness: $(q_{it} = open_{it})$	
Threshold value $(\hat{\theta}_{open})$	0.318
F-test	63.902
Bootstrapped p-values	(0.260)
Exchange rate volatility: $(q_{it} = \sigma_{it}^{\Delta e})$	
Threshold value $(\hat{\theta}_{\sigma^{\Delta e}})$	0.082
F-test	78.738
Bootstrapped p-values	(0.010)

Note: Table reports threshold estimates ($\hat{\theta}$), F-test of the null hypothesis of no threshold effect and bootstrapped p-values (obtained from 1000 bootstrap replications).

[11] To obtain exchange rate volatility series, we start by computing the standard deviation of changes in exchange rate for the first quarter 1994:1 during the last five years and, then, we slid forward this window quarter by quarter throughout our estimation period (1994-2010).

Table 8: Country Classification

Inflation Regime			**Degree of Openness**			**Exchange Rate Volatility**		
Low Inflation	**Moderate inflation**		**Less Open**	**More Open**		**Low Volatility**	**High Volatility**	
Austria	Australia	Portugal	Australia	Austria	Netherlands	Australia	Belguim	Poland
Belguim	Czech Republic	Slovak Republic	Germany	Belguim	Poland	Austria	Czech Republic	Portugal
Canda	Denmark	United States	Spain	Canda	Portugal	Canda	Germany	Slovak Republic
Switzerland	Spain		France	Switzerland	Slovak Republic	Switzerland	Spain	Sweden
Germany	Greece		United Kingdom	Czech Republic	Sweden	Denmark	Finland	
Finland	Ireland		Italy	Denmark		France	Ireland	
France	Iceland		Japan	Finland		United Kingdom	Iceland	
United Kingdom	Italy		Norway	Greece		Greece	Italy	
Japan	Korea		New Zealand	Ireland		Japan	Korea	
Luxembourg	Netherlands		United States	Iceland		Norway	Luxembourg	
New Zealand	Norway			Korea		United States	Netherlands	
Sweden	Poland			Luxembourg			New Zealand	
12 countries	**15 countries**		**10 countries**	**17 countries**		**11 countries**	**16 countries**	

Note: Last line denote number of countries in each class. The volatility of the exchange rate changes, $\sigma^{\Delta e}$, is computed as the standard deviation of quarterly percentage changes in the exchange rate.

6.3 Regime dependence of ERPT

Following countries classification, now we must perform estimation for each panel group of countries. So before applying FM-OLS and DOLS estimators, we proceed by testing panel unit root for individual series within each group (high and low inflation, more and less open countries, and more and less exchange rate volatility). Results from IPS tests (reported in Appendix B.1) show that most of variables are $I(1)$. Then, we provide the presence of cointegration relationship by using Pedroni cointegration tests for different sub-sample panel of countries (Appendix B.2). Almost all of tests lead us to reject the null of non-cointegration.

Estimates of long-run ERPT for each group of countries reported in Table 9. We begin with the inflation rate as a macro-determinant of the extent of pass-through. In view of results, low inflation countries experience long run import prices elasticity equal to 0.53% by FM-OLS. While one percent exchange rate depreciation causes an increase in import prices by 0.75% in high inflation countries. Result remain robust when using DOLS method. Thus, ERPT is found to be higher in high inflation environment countries. It is evident that this finding corroborates the convention wisdom of the positive link between Inflation and pass-through (Taylor (2000)). That is, countries with higher rates of inflation should have higher rates of pass-through of exchange rates into import prices. Our results provide an evidence of regime-dependence of ERPT with respect to inflation environment and this latter would be an important source of heterogeneity in pass-through across countries.

For our second macro-determinant, i.e. import share, one can expect a positive connection between openness and pass-through: the more a country is open, the more import prices respond to exchange rate fluctuations. According to our results this positive link seems to be weak. Both FM-OLS and DOLS show a long-run ERPT of roughly 0.56% in less open economies, which is little smaller than in the more open ones (0.68% by FM-OLS). The 95% confidence band shows that the extent of pass-through seems to do not differ strongly between the two group of country, especialy according to DOLS estimators. As we mentioned above, there is no conclusive empirical results in the literature about the relevance of degree of openness. For nine developed countries, McCarthy (2007) shows that association is not significant between import share and pass-through[12]. However, Barhoumi (2006) found a positive correlation of pass-through-openness in panel cointegration framework. The main difference with our analysis is that the measure of openness used in Barhoumi (2006) is the tariffs barriers. The author found that lower tariff barriers countries experience a higher long run pass-through than higher tariff barriers.

[12]Similarly, Choudhri & Hakura (2006) found a little evidence of a positive relationship between ERPT to consumer prices and openness

Table 9: Long run Pass-Through Estimates for different country regime

	Inflation Regime		**Degree of openness**		**Exchange rate volatility**	
	Low Inflation	**High Inflation**	**Less Open**	**More Open**	**Less volatile**	**More volatile**
FMOLS	0,53**	0,75**	0,57**	0,68**	0,47**	0,79**
	[0,49 ǀ 0,57]	[0,70 ǀ 0,81]	[0,53 ǀ 0,60]	[0,62 ǀ 0,57]	[0,43 ǀ 0,52]	[0,74 ǀ 0,84]
DOLS	0,51**	0,82**	0,56**	0,66**	0,39**	0,74**
	[0,46 ǀ 0,55]	[0,76 ǀ 0,89]	[0,52 ǀ 0,60]	[0,58 ǀ 0,75]	[0,35 ǀ 0,43]	[0,69 ǀ 0,79]
	12 countries	15 countries	10 countries	17 countries	11 countries	16 countries

Note: ** indicate statistical significance at the 5 percent level. 95% confidence intervals are reported between square brackets.

Finally, we raise the question about the relevance of exchange rate volatility in explaining the long run pass-through. In fact, it is expected that import prices responsiveness would be higher when volatility of exchange rate is larger. As pointed by Devereux & Engel (2002), the relative stability of market destination currency plays a substantial role in determining pass-through. Countries with low relative exchange rate variability would have their currencies chosen for transaction invoicing. Thereby, local-currency pricing (LCP) would prevailing and pass-through is less than complete. In view of our results, pass-through elasticity is about 0.40% in less volatility exchange rate countries, but import prices increase by 0.74% following one percent nominal depreciation in high volatility countries (according to DOLS estimates). There is significant difference between the two groups, and results are robust across FM-OLS and DOLS estimates. Empirically, this finding is consistent with Campa & Goldberg (2005) for whom higher home currency volatility is significantly associated with lower ERPT.

It is important to mention that this positive link between is not as obvious as one would think. In his VAR Study, McCarthy (2007) suggest that that pass-through should be less in countries where the exchange rate has been more volatile. The author argued that greater home currency volatility may make exporters more willing to adjust profit margins, which reduces measured pass-through. In his panel of developing countries, Barhoumi (2006) gives support to this intuition. He obtains a lower pass-through for fixed exchange rate regime countries which are defined as panel group with less volatile exchange rate.

7 Conclusion

This paper has examined the long run exchange rate pass-through (ERPT) into import prices using panel cointegration approach. We first provide a strong evidence of incomplete ERPT in sample of 27 OECD countries. On the long run, import prices do not move one to one following exchange rate depreciation. Both FM-OLS and DOLS estimators show that pass-through elasticity does not exceed 0.70%. These results are in line with estimates in the literature of exchange rate pass-through into import prices for industrialized countries. When considering individual estimates for our panel of 27 countries, we can note a cross-country difference in the long run ERPT, with the highest import prices reactions are recorded in Poland by 0.98% followed by Czech Republic with 0.95%. It is important to mention that there is an evidence of complete pass-through for 5 out of 27 countries, namely Czech Republic, Italy, Korea, Luxembourg and Poland. The cross-county differences in the pass-through lead us to the question of what are the underlying determinants of pass-through. Then, when split our sample in two inflation country regime, we find that high inflation countries have experienced a higher degree of ERPT than lower inflation ones. These findings are in line with Taylor's hypothesis. Another potential source of cross-country differences is home currency volatility. In view of our results, import prices responsiveness would be lower in countries with less volatile exchange rate. This can be explained by foreign firms' behaviors which are willing to set their prices in stable currency country (local currency pricing (LCP)). We can mention that we find a weak evidence of a positive link between degree of openness and ERPT which is commonly agreed in the pass-through literature.

Appendices

A Estimation methods

A.1 FM-OLS Mean Group Panel Estimator: Pedroni (2001)

We consider the following fixed effect panel cointegrated system:

$$y_{it} = \alpha_i + x_{it}^{'}\beta + \varepsilon_{it}, \quad t = 1, \ldots T, \tag{16}$$

$x_{it}^{'}$, can in general be a m dimensional vector of regressors which are integrated of order one, that is:

$$x_{it} = +x_{it-1} + u_{it}, , \forall i \tag{17}$$

Where the vector error process $\xi_{it} = (\varepsilon_{it}, u_{it})^{'}$ is stationary with asymptotic covariance matrix:

$$\Omega_{it} = \lim_{T\to\infty} E\left[T^{-1}\left(\sum_{t=1}^{T}\xi_{it}\right)\left(\sum_{t=1}^{T}\xi_{it}^{'}\right)\right] = \Omega_i^0 + \Gamma_i + \Gamma_i^{'}. \tag{18}$$

Ω_i^0, is the contemporaneous covariance and, Γ_i, is a weighted sum of autocovariances.

The long run covariance matrix is constructed as follow: $\begin{bmatrix}\Omega_{11i} & \Omega_{21i}^{'} \\ \Omega_{21i} & \Omega_{22i}\end{bmatrix}$, where, Ω_{11i}, is the scalar long run variance of the residual, ε_{it}, and, Ω_{22i}, is the long run covariance among the, u_{it}, and, Ω_{21i}, is vector that gives the long run covariance between the residual, ε_{it}, and each of the u_{it}.

For simplicity, we will refer to, x_{it}, as univariate. So according to Pedroni (2001), the expression for the group-mean panel FM-OLS estimator (for the between dimension) is given as:

$$\hat{\beta}_{GFM} = N^{-1}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)^2\right)^{-1} \times \left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)\, y_{it}^{*} - T\hat{\gamma}_i\right) \tag{19}$$

Where $y_{it}^{*} = (y_{it} - \bar{y}_i) - \frac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}} \Delta x_{it}$, and $\hat{\gamma}_i \equiv \hat{\Gamma}_{21i} - \Omega_{21i}^{0} - \frac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}} \left(\hat{\Gamma}_{22i} - \Omega_{22i}^{0} \right)$, with $y_i = \frac{1}{T} \sum_{t=1}^{T} y_{it}$ and $x_i = \frac{1}{T} \sum_{t=1}^{T} x_{it}$ refer to the individual specific means.

The Pedroni between FM-OLS estimator, $\hat{\beta}_{GFM}$, is the average of the FMOLS estimator computed for each individual, $\hat{\beta}_{FM,i}$, that is:

$$\hat{\beta}_{GFM} = N^{-1} \sum_{i=1}^{N} \hat{\beta}_{FM,i} \tag{20}$$

The associated t-statistic for the between-dimension estimator can be constructed as the average of the t-statistic computed for each individuals of the panel:

$$t_{\hat{\beta}_{GFM}} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{\beta}_{FM,i}} \tag{21}$$

Where $t_{\hat{\beta}_{FM,i}} = \left(\hat{\beta}_{FM,i} - \beta_0 \right) \left(\hat{\Omega}_{11i}^{-1} \sum_{t=1}^{T} (x_{it} - \bar{x}_i)^2 \right)^{1/2}$.

A.2 DOLS Mean Group Panel Estimator: Pedroni (2001)

The DOLS regression can be employed by augmenting the cointegrating regression with lead and lagged differences of the regressors to control for endogenous feedback effects. Thus, we can obtain from the following regression:

$$y_{it} = \alpha_i + \beta_i x_{it} + \sum_{k=-K_i}^{K_i} \gamma_{it} \Delta x_{it-k} + \varepsilon_{it}, \tag{22}$$

The group-mean panel DOLS estimator is construct as:

$$\hat{\beta}_{GD} = N^{-1} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} Z_{it} Z_{it}^{\prime} \right)^{-1} \left(\sum_{t=1}^{T} Z_{it} \tilde{y}_i \right) \tag{23}$$

Where $Z_{it} = (x_{it} - \bar{x}_i, \Delta x_{it-K}, ..., \Delta x_{it-K})$ is a the $2(K+1) \times 1$ vector of regressors and $\tilde{y}_{it} = y_{it} - \bar{y}_i$.

The DOLS estimator for the i^{th} member of the panel is written as:

$$\hat{\beta}_{D,i} = \left(\sum_{t=1}^{T} Z_{it} Z_{it}^{'}\right)^{-1} \left(\sum_{t=1}^{T} Z_{it} \tilde{y}_{i}\right) \tag{24}$$

So that the between-dimension estimator can be constructed as

$$\hat{\beta}_{GD} = N^{-1} \sum_{i=1}^{N} \hat{\beta}_{D,i} \tag{25}$$

If the long-run variance of the residuals from the DOLS regression (23) is:

$$\sigma_i^2 = \lim_{T\to\infty} E\left[T^{-1}\left(\sum_{t=1}^{T} \varepsilon_{it}\right)^2\right] \tag{26}$$

According to Pedroni, the associated t-statistic for the between-dimension estimator can be constructed as:

$$t_{\hat{\beta}_{GD}} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{\beta}_{D,i}} \tag{27}$$

Where $t_{\hat{\beta}_{D,i}} = \left(\hat{\beta}_{D,i} - \beta_0\right)\left(\hat{\sigma}_i^{-2} \sum_{t=1}^{T} (x_{it} - \bar{x}_i)^2\right)^{1/2}$.

A.3 Estimation of Panel Single Threshold Model: Hansen (1999)

Equation (15) in the text can be written as follows:

$$y_{it} = \alpha_i + \beta^{'} x_{it}(\theta) + \varepsilon_{it}, \tag{28}$$

Where y_{it} is the dependent variable, $x_{it}(\theta) = \begin{pmatrix} x_{it} I(q_{it} \leq \theta) \\ x_{it} I(q_{it} > \theta) \end{pmatrix}$ is a k-dimensional vector of exogenous variables and $\beta = (\beta_1^{'}, \beta_2^{'})$.

After removing the individual-specific means, α_i, using the within transformation estimation techniques, the OLS estimator of β is given by:

$$\hat{\beta}(\theta) = (X^*(\theta)^{'} X^*(\theta))^{-1} X^*(\theta)^{'} Y^* \tag{29}$$

Where X^* and Y^* denote the stacked data over all individuals after removing the individual specific means.

The vector of regression residuals is $\hat{\varepsilon}^*(\theta) = Y^* - X^*(\theta)\hat{\beta}(\theta)$ and the sum of squared errors can be written as

$$S_1(\theta) = \hat{\varepsilon}^*(\theta)^{'} \hat{\varepsilon}^*(\theta) = Y^{*'} \left(I - X^*(\theta)^{'} (X^*(\theta)^{'} X^*(\theta))^{-1} X^*(\theta)^{'}\right) Y^* \tag{30}$$

In a second step Hansen (1999) recommend the estimation of the threshold θ by least squares which is achieved by minimization of the concentrated sum of squared errors $S_1(\theta)$. Then, the least squares estimators of $\hat{\theta}$ is given by

$$\hat{\theta} = \underset{\theta}{\operatorname{argmin}} \, S_1(\theta) \tag{31}$$

Hence, the resulting estimate for the slope coefficient is obtained by $\hat{\beta} = \hat{\beta}(\hat{\theta})$. The residual vector is $\hat{\varepsilon}^* = \hat{\varepsilon}^*(\hat{\theta})$ and residual variance is defined as:

$$\hat{\sigma}^2 = \frac{1}{N(T-1)} \hat{\varepsilon}^{*'} \hat{\varepsilon}^* = \frac{1}{N(T-1)} S_1(\hat{\theta}) \tag{32}$$

B Stationarity and cointegration tests for different regimes

B.1 Panel unit root tests

Table 10: IPS tests for different country regime

	Level		**First difference**	
	Intercept	**Intercept & trend**	**Intercept**	**Intercept & trend**
Low Inflation				
mp_{it}	-0.1405	-1.0594	-7.9740	-9.1933
e_{it}	-0.8485	-0.8553	-10.9606	-8.0266
y_{it}	-0.1405	3.5652	-9.3488	-8.7950
w^*_{it}	0.3445	-0.2818	-8.8177	-9.5902
High Inflation				
mp_{it}	-0.6878	-0.2549	-5.9312	-8.2145
e_{it}	2.2388	0.0321	-7.6749	-7.8984
y_{it}	-0.5381	3.1936	-6.7155	-6.6649
w^*_{it}	0.2137	-1.1218	-5.3517	-6.7586
Low Openness				
mp_{it}	-0.6883	-1.0806	-4.9044	-4.9137
e_{it}	0.6127	-1.6987	-7.0398	-9.5322
y_{it}	-1.4890	-0.5947	-3.8447	-3.7240
w^*_{it}	2.2661	-0.8751	-3.9590	-3.9204
High Openness				
mp_{it}	0.0553	-0.2854	-3.7535	-3.7441
e_{it}	2.6988	0.6116	-6.2117	-6.7251
y_{it}	0.1784	1.2138	-6.2556	-5.7448
w^*_{it}	0.5523	-0.3322	-6.5015	-3.2179
Low Volatility				
mp_{it}	0.1393	-0.5981	-5.1119	-5.7440
e_{it}	1.4496	0.3933	-5.3775	-4.3244
y_{it}	-1.7040	4.0617	-0.4306	-8.9381
w^*_{it}	1.4477	-0.9389	-6.5228	-3.9717
High Volatility				
mp_{it}	-1.0527	-0.2813	-3.5419	-5.6684
e_{it}	0.4843	-0.7306	-2.9523	-7.5928
y_{it}	0.8293	2.0367	-3.3253	-6.4672
w^*_{it}	1.5506	-0.2651	-4.9772	-7.2921

Note: For the IPS tests, the critical value at the 5% level is -1.81 for model with an intercept and -2.44 for model with intercept and linear time trend. Individual lag lengths are based on Akaike Information Criteria (AIC).

B.2 Panel cointegration

Table 11: Pedroni tests for different countries regimes

	Inflation		Openness		Exchange rate volatility	
	Low	**High**	**Low**	**High**	**Low**	**High**
panel v-stat	5.847	3.812	6.145	3.406	5.160	4.715
panel rho-stat	-5.339	-3.273	-5.297	-3.289	-3.669	-5.564
panel pp-stat	-5.746	-3.400	-5.527	-3.640	-4.186	-5.402
panel adf-stat	-5.114	-1.770	-5.608	-0.721	-2.692	-4.489
group rho-stat	-4.060	-3.229	-4.578	-2.509	-2.770	-4.911
group pp-stat	-5.520	-3.704	-5.641	-3.513	-4.080	-5.582
group adf-stat	-4.554	-1.832	-5.803	0.015	-2.702	-4.247

Note: 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).

C Threshold levels according to sum of squared residuals

Figure 1: Sum of squared residuals function when $q_{it} = \pi_{it}$

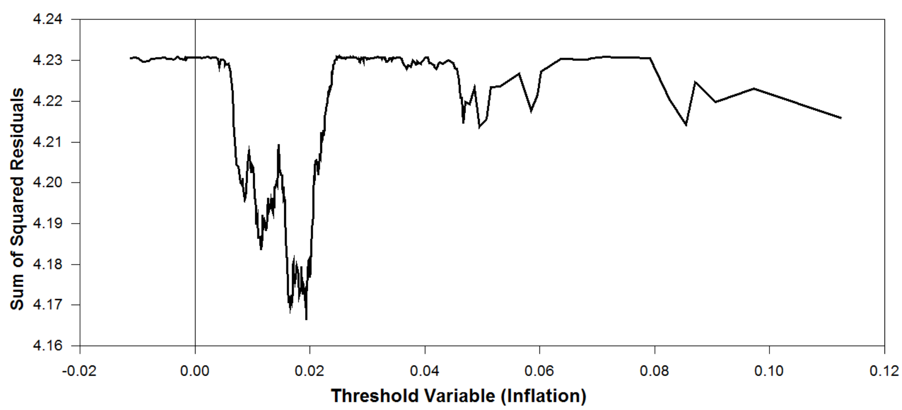


Figure 2: Sum of squared residuals function when $q_{it} = open_{it}$

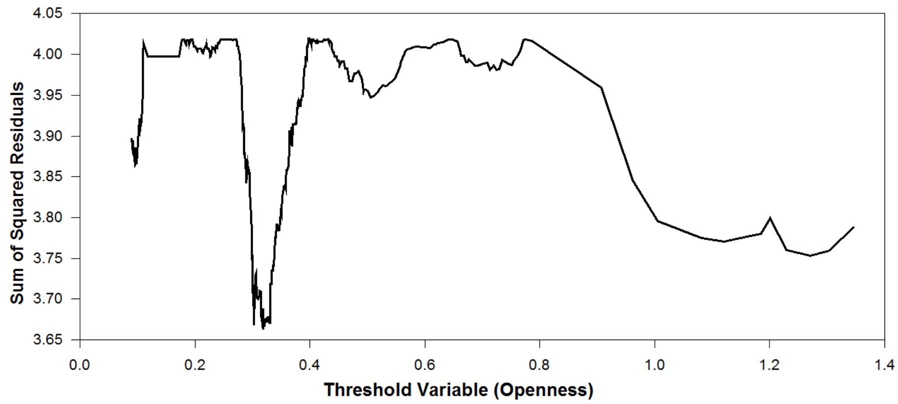
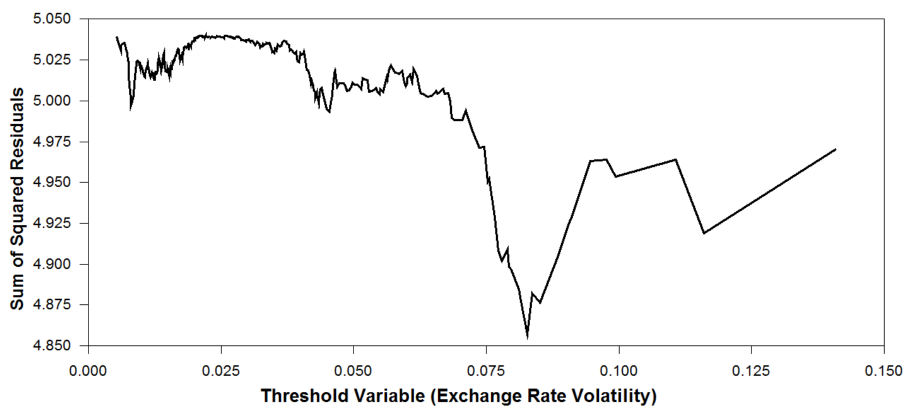


Figure 3: Sum of squared residuals function when $q_{it} = \sigma_{it}^{\Delta e}$



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