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Nidhaleddine Ben Cheikh[1]

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JEL : C23, E31, F31, F40
Keywords: Exchange Rate Pass-Through, Import Prices, Nonstationary Panel data

Author

[1] CREM, University of Rennes 1, 7 Place Hoche, 35065, Rennes Cedex, France
Tel.: +33 223 23 35 48. E-mail address: nbeneche@univ-rennes1.fr

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*Tel.: +33 223 23 35 48. *E-mail address*: nbeneche@univ-rennes1.fr.

1 Introduction

The issue of exchange rate pass-through (ERPT) to domestic prices has long been of interest in debates about the international transmission of monetary shocks and the optimal choice of exchange rate and monetary policy regimes. As ERPT is a channel linking exchange rates with prices, it is usually considered as one of the key determinants of monetary policy design. ERPT is traditionally defined as the degree of sensitivity of import prices to a one percent change in exchange rates in the importing nation's currency. However, it is commonly argued in pass-through literature that the import prices do not move one-to-one following exchange variations. In other words, ERPT is found to be incomplete, and this has been observed in several industrial countries. In fact, there is decline in pass-through since the early 1990's. Although there is no consensus about the conditions that lead to a low pass-through, it is generally agreed that moving towards more stable inflation environment has played an important role in the recent fall in ERPT.

The 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. It is important to mention that others macroeconomic determinants are often mentioned in pass-through literature, such as degree of openness and exchange rate volatility. It is well-known that ERPT comprising two stages process. First stage concerning the exchange rate transmission to import prices, and the second stage deals with consumer prices adjustments to exchange rate movements. So, literature is divided with respect to these two stages. In our empirical, we look only at the ERPT to import prices. More precisely, we consider the pass-through from exchange rate to import prices for a long-run time horizon, since a cointegrated equilibrium relationship could be found in pass-through equation.

To address this issue, we use a panel cointegration approach in order to give a relevant definition to the long measure of the pass-through. We can note that there has been an increasing use of unit root and cointegration analysis in the context of panel data. Therefore, in our empirical analysis we follow Pedroni (2001) methodology by applying FMOLS and DOLS group mean estimators to obtain long run ERPT for 27 OECD countries. Little is about long run pass-through in this context, and the aim of our paper is to fill this gap and to provide new evidence by using these recent panel data techniques. Another important issue is to explain the cross-country differences in the import prices responsiveness to exchange rate movements. In our analysis, we explore three macroeconomic factors, i.e. inflation rate, degree of openness and exchange rate volatility which are potential sources of heterogeneity in ERPT. Due to the important implications of incomplete pass-through for monetary union, in the final part of our analysis, we focus on the case of the euro area by taking a sub-sample of 12 European countries. Our goal is to assess the behavior of ERPT since the collapse of Breton-Woods era and try to relate it to the change in the inflation environment.

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 the main macroeconomic factors determining ERPT. In Section 7, we focus on The EMU countries by assessing how pass-through has changed over time. Section 8 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. Especially, this finding remains strong 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, micro-determinants like pricing strategies of firms is one of major reason of partial ERPT.

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 (henceforth PTM[1]) strategy by setting different prices for different destination markets. 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

[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.

following destination market currency depreciation, PTM occurs and pass-through to import prices is less than complete. In the extreme case, prices do not to vary in the currency of importing country, this is refers to local currency pricing (LCP) strategy, then pass-through would be equal to zero.

In a more recent literature, there has been a growing interest in examining the relationship of ERPT and macroeconomic factors. One of the most convincing factors is the inflation environment in each country. This latter macro-determinant is brought by Taylor (2000) who argues that the responsiveness of prices to exchange rate fluctuations depends positively on inflation. So pass-through 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. It is worth noting that many empirical studies gave a supportive evidence to the Taylor's hypothesis, such as Choudhri & Hakura (2006), Gagnon & Ihrig (2004) and Bailliu & Fujii (2004).

Another important macroeconomic determinant of pass-through is the exchange rate volatility. This latter would be positively associated with higher import price pass-through. Most of pass-through studies find that countries with low nominal exchange rate volatility have a lower ERPT. 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 pass-through is less than complete. Empirically, Campa & Goldberg (2005) find that exchange rate volatility is positively associated with higher import price pass-through in 23 OECD countries, although microeconomic factors play a much more important role in determining the pass-through. 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. Several Studies have tested the relevance of others macro-determinant, especially, 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 little evidence of a positive relationship between openness and pass-through to import price.

In our empirical, we focus on the ERPT 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 the 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 time series and nonstationary panel data techniques, their study propose to restore the cointegrated long run equilibrium in the ERPT relationship between the variables in levels. 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, i.e. the gain of statistical power of testing procedure to identify non-spurious cointegration between the variables[2].

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 summarize 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 somecountries of the European Union. Nevertheless, our sample is larger since we consider 27 OECD in the first part of our analysis. 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. Furthermore, we can mention that De Bandt *et al.* (2007) measure the long run ERPT at the disaggregate level for 11 of the euro area countries, while we use aggregate import prices data for nearly the same sample of country.

[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 a Panel Cointegration analysis**

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. Countries with fixed exchange rate, lower tariff barriers and higher inflation regimes exhibit a higher long-run ERPT.
Holmes (2006)	Monthly data (1972:4- 2004:6) for 12 European Union countries	Employment of panel data cointegrating techniques. Using dynamic ordinary least squares (DOLS) which provides group mean estimates of long-run ERPT to consumer prices.	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 (progression towards and introduction of the single currency in the 1990s).
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 (Pedroni (1999)) with no breaks and Banerjee & Carrion-i Silvestre (2006) with. estimated breakpoint). -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.	A panel cointegration approach. FMOLS (Pedroni (2001)) 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 nominal effective exchange rate over time. This declining in the long-run pass-through is accompanied by decreasing in inflation rates occurring since the mid-1990s.

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 (PCP is prevailing). And if $\Phi = -1$, exporters decide not to vary the prices in the destination country currency, 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 to 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, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom and United States. The data are quarterly and span the period 1994-2009. 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:

$$W^*_{it} = Q_{it} \times \frac{W_{it}}{E_{it}} \tag{10}$$

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:

$$w_{it}^{*} = q_{it} + e_{it} - p_{it} \tag{11}$$

Since the 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.

4 Empirical methodology

4.1 Panel unit root tests for dynamic heterogeneous panels

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[4],

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

[4]In our case all variables are assigned to y_{it}.

$$y_{it} = \mu_i + \rho_i y_{it-1} + \sum_{j=1}^{p_i} \varphi_{it} \Delta y_{it-j} + \gamma_i t + \varepsilon_{it}, \qquad t = 1, ...T, \tag{12}$$

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:

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^{N} t_{\rho i}, \tag{13}$$

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[5].

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

Variables	**Level**	**First difference**	**Level**	**First difference**
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
mp_{it}	-0,5301	-1,3952	-14,0574	-17,2283
e_{it}	0,6973	-0,144	-10,3162	-11,1541
y_{it}	-1,1128	1,078	-7,577	-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).

4.2 Tests for panel cointégration

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

[5] 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 $T = 70$.

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:

$$\begin{aligned} &y_{it} = \alpha_i + \delta_{it} + \beta_{1i}x_{1it} + \beta_{2i}x_{2it} + ... + \beta_{Ki}x_{Kit} + \varepsilon_{it}, \\ &\qquad i = 1, ..., N, \ t = 1, ...T, \ k = 1, ..., K \end{aligned} \tag{14}$$

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:

$$\frac{\chi_{NT} - \mu\sqrt{N}}{\sqrt{\upsilon}} \rightarrow N(0, 1), \tag{15}$$

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	**1994Q1 - 2010Q1**
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 exchange rate pass-through estimations

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[6]. 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[7]. 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[8]. To check robustness of our result, we also reporting estimation results for the pooled OLS and 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 (according to t-statistics reported between square brackets in Table 4). 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

[6]Brief details of these methods are available in Appendix A.

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

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

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, then a partial import prices it is expectable for OECD countries.

Table 4: **Panel Estimates For 27 OECD countries over 1994q1-2009q4**

	Dependent Variable: Import Price Index			
	Group mean FM-OLS	**Group mean DOLS**	**Fixed effects**	**Pooled OLS**
e_{it}	0.67***	0.69***	0.70***	0.71***
	(30.21)	(26.69)	(33.01)	(31.97)
	[16.71]	[16.89]	[10.29]	[12.60]
y_{it}	0.27***	0.20***	0.23***	0.016***
	(6.15)	(6.40)	(11.86)	(17.63)
w^*_{it}	0.68***	0.71***	0.214***	0.28***
	(7.09)	(6.89)	(8.215)	(18.40)

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. The t-statistics in square brackets are based on the null of unitary elasticity.

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)[9]. 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 pricing-to-market (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 show that the highest import prices reaction is in Poland by 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

[9]For exemple, see Smets and Wouters (2002).

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 lower long run import prices adjustment in the United States by 0.38%, which is a common result in the literature (Campa & Goldberg (2005)) find 41% ERPT elasticity).

Table 5: **Long run individual Pass-Through for 27 OECD Countries**

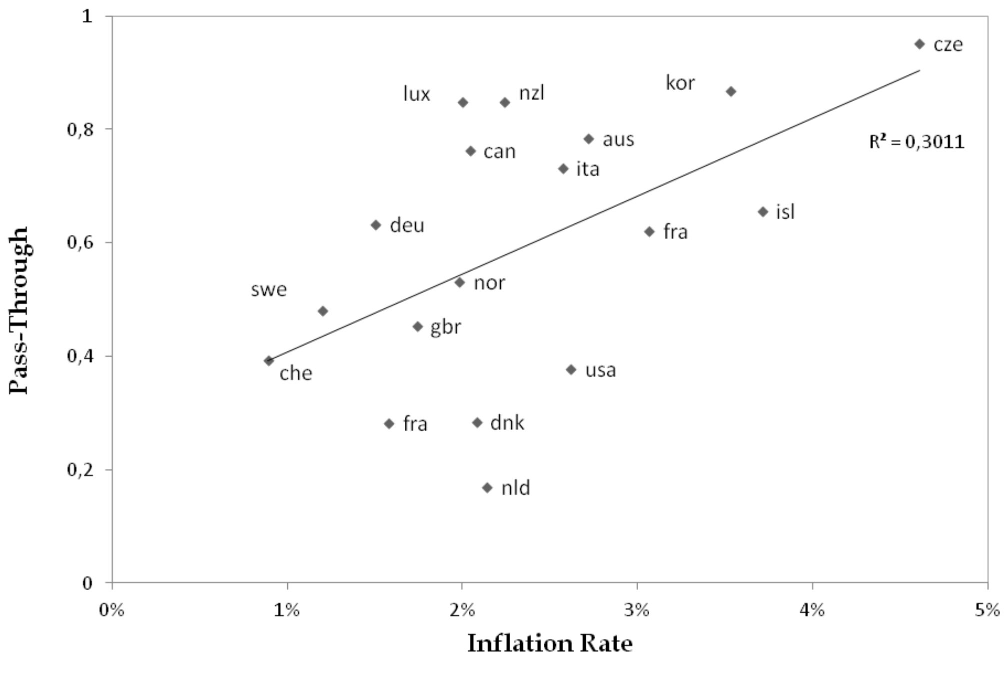
	Results from FM-OLS method		
Country	**FM-OLS**	*t-stat for* $H_0 : \beta = 0$	*t-stat for* $H_0 : \beta = 1$
Australia	0,78*	32,7	9,04
Austria	-0,08	-0,23	3,28
Belguim	-0,04	-0,28	6,57
Canda	0,76*	18,42	5,75
Switzerland	0,39*	3,32	5,14
Czech Republic	$0,95^{*\#}$	10,75	0,54
Germany	0,63*	4,2	2,44
Denmark	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 $H_0 : \beta = 1$.

Having estimated long run ERPT coefficients, we next examine whether in line with Taylor's hypothesis there is evidence of a positive correlation between pass-through and inflation. The idea is exporters pricing strategies may be endogenous to a country's relative monetary stability. So for more stable inflation destination countries, foreign firms are willing to adopt local currency price stability (LCP) and pass-through would be incomplete. To obtain some insights on this potential positive link, we calculate year-on-year quarterly inflation rates and take the mean values over the period 1994-2009. These statistics for our 27 OECD countries are reported in Table 6. We should note that Japan has the lowest inflation rate with a negative value (-0.1%), while Poland experiences the highest rate exceeding 8 percent. So, in order to establish a relevant ERPT-Inflation correlation, we eliminate Japan and Poland from analysis and also countries with non-significant pass-through.

By visual inspection of Figure 1, we can note a positive slope arising from ERPT-Inflation. This is a strong evidence of a positive and significant association between the pass-through and the average inflation rate across countries. This finding appears overall supportive to Taylor's hypothesis. Countries with high inflation environment would experience a higher degree of pass-through. According to Campa & Goldberg (2005), although macroeconomic variables play limited role in explaining cross-country differences in ERPT, inflation rates affect significantly the extent to which exchange rate changes are "passed through" import prices.

Figure 1: ERPT and Inflation Correlation



Sources: Personal Calculation.

6 Macroeconomic Factors Affecting Pass-Through

Cross-country differences in the long run import prices adjustment to exchange rate would raise the question of what are the underlying determinants of pass-through. In the previous section, we have shown an important determinant of ERPT, i.e. inflation rate. Many empirical analyses have explored the influence of other macroeconomic variables such as, Exchange rate volatility and degree of openness. To pursue explanation of sources of this long run heterogeneity, we now examine some macroeconomic factors that may affect pass-through.

Three main factors are selected for this purpose: inflation rates to give further evidence in support of Taylor hypothesis; degree of openness measured by the mean of imports as a percentage of domestic demand over 1994-2009; and exchange rate volatility changes, $\sigma_{\Delta e}$ as proxied by the standard deviation of quarterly percentage changes in the exchange rate. A summary of the average macroeconomic conditions in our country sample over 1994-2009 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 to 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 (1994-2009)**

Country	Mean Rate of Inflation (%)	Openness (%)	Exchange Rate Volatility (%)
Australia	2,7	16,3	8,6
Austria	1,8	44,4	9,2
Belguim	1,8	74,7	9,6
Canda	2	39,1	5,2
Switzerland	0,9	40,8	8,4
Czech Republic	4,6	75	10
Germany	1,5	34,3	9,2
Denmark	2,1	43,5	8,8
Spain	3,1	32,4	11,3
Finland	1,4	35	13,6
France	1,6	27,4	8,8
United Kingdom	1,7	26,6	7,7
Greece	4,3	33,5	8
Ireland	3,7	38,6	8,9
Iceland	3,2	67,5	15,1
Italy	2,6	25,5	11,1
Japan	-0,1	9,8	8,9
Korea	3,5	33	13,8
Luxembourg	2	135,2	9,6
Netherlands	2,1	64,7	9,3
Norway	2,2	32	8,1
New Zealand	2	26	10,8
Poland	8,4	34,1	15,3
Portugal	3	35,5	9,8
Slovak Republic	6,7	77,2	10,5
Sweden	1,2	38,4	11,5
United States	2,6	14,1	5,8
Average	**2,7**	**42,8**	**9,9**

Note: The volatility of the exchange rate changes,$\sigma_{\Delta e}$, is computed as the standard deviation of quarterly percentage changes in the nominal effective exchange rate.

Concerning the inflation rate criteria, we can note that Choudhri & Hakura (2006) classify their 71 countries into three groups based on the average of inflation rate. 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. For our 27 countries, we choose to decompose them into tow country groups: a low inflation regime and a high inflation regime. According to Table 6, the mean of inflation rate in our 27 OECD countries is close to 3 percent. So, we can define "low inflation" countries with less than 3% of inflation rate and "high inflation" as group of countries with more than 3 percent. To make sure that our classification is not rather arbitrary, we experiment Hansen (1999) threshold regression for panel data which permit us to divide our sample of into different classes of countries based on their inflation regime[10]. This methodology allows us to select the threshold level by estimating this parameter from the sample. Another important consideration is that threshold variable must be exogenous, therefore we assume that is the case for inflation rate[11]. The single threshold model can be written as follow:

The single threshold model can be written as follow:

$$mp_{it} = \alpha_i + \beta_1 e_{it} I(\pi_{it} \leq \theta) + \beta_2 e_{it} I(\pi_{it} \geq \theta) + \gamma y_{it} + \delta w^*_{it} + \varepsilon_{it}, \tag{16}$$

Where inflation rate, π_{it}, correspond to the threshold variable. Therefore, the observations are divided into two regimes depending on whether, π_{it}, is smaller or larger than the threshold, θ. We allow only for the pass-through elasticity to switch between regimes so each one is distinguished by a different slop, respectively, β_1 and β_1. We then estimate the equation (14) by least square assuming sequentially zero and one single threshold. Estimation result provides a 0.028 as a threshold inflation level with F-stat for a single threshold is highly significant (168.03). It is evident that using Hansen (1999) approach gives support evidence on our sample splits: threshold inflation (2.8%) is roughly equal the mean rate of inflation in our sample (2.7%). Thus, as we mentioned above, countries characterized by mean inflation rate less than 3% will be considered as low inflation countries, while countries having mean inflation larger than 3% will be considered as high inflation countries. This gives us 17 low inflation countries and 10 countries with mean inflation rate more than 3% (see Table 7).

[10]Since we divide our sample into two groups, we assume a single threshold in our model. Hansen (1999) methodology can be used for multiple thresholds.

[11]This is a strong assumption because one can think that inflation rate can be influenced by the extent of pass-through.

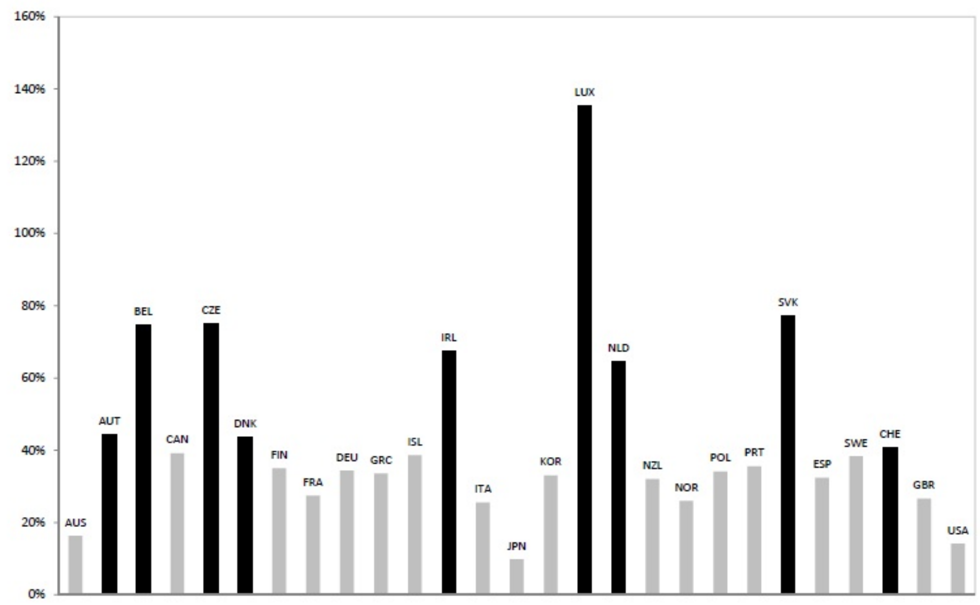
Table 7: **Country Classification**

Inflation Regime			**Degree of openness**			**Exchange rate volatility**		
Low Inflation		**High Inflation**	**Less Open**		**More Open**	**Less volatile**		**More volatile**
Australia	Netherlands	Czech Republic	Australia	Korea	Austria	Australia	Luxembourg	Czech Republic
Austria	New Zealand	Greece	Canada	New Zealand	Belgium	Austria	Netherlands	Finland
Belgium	Norway	Iceland	Finland	Norway	Czech Republic	Belgium	Norway	Iceland
Canada	Sweden	Ireland	France	Poland	Denmark	Canada	Portugal	Italy
Denmark	Switzerland	Italy	Germany	Portugal	Ireland	Denmark	Switzerland	Korea
Finland	United Kingdom	Korea	Greece	Spain	Luxembourg	Germany	United Kingdom	New Zealand
France	United States	Poland	Iceland	Sweden	Netherlands	France	United States	Poland
Germany		Portugal	Italy	United Kingdom	Slovak Republic	Greece		Slovak Republic
Japan		Slovak Republic	Japan	United States	Switzerland	Ireland		Spain
Luxembourg		Spain				Japan		Sweden
17 countries		**10 countries**	**18 countries**		**9 countries**	**17 countries**		**10 countries**

Note: Last line denote number of countries in each class. Italy is added to high inflation country, since it has a high inflation level in the 1990's (more than 4 percent). The volatility of the exchange rate changes, $\sigma_{\Delta e}$, is computed as the standard deviation of monthly percentage changes in the exchange rate. In our 27 OECD countries sample the volatility average is about 10%, so we define "more volatile" countries with exchange rate fluctuations exceeding 10% and "less volatile" ones with fluctuations lower than 10%.

Moving to the second macroeconomic factor, i.e. Import share, we notice that the mean value of degree of openness is about 43%. Table 6 and Figure 2 give the import shares of our 27 OECD countries. Knowing that this variable don't move considerably over our sample, we then define "more open countries" as group with more than 40% of imports (as percentage of Domestic demand) and "less open countries" with degree of openness less than 40%. Our choice of 40% instead of 43% as threshold level is to include more countries in the "more open" group, and thus to have enough observation to estimate this sub-sample panel. In all, we obtain 18 less open countries and 9 countries with import share exceeding 40% (see Table 7).

Figure 2: Share of Imports (as a percentage of domestic demand over 1995-2009)



Sources: OCDE

Finally, the last criterion which is can explain differences in pass-through across countries 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 quarterly 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, $\sigma_{\Delta e}$, over 1994-2009. According to Table 6, the average of this macroeconomic factor is about 10% in our county sample. So, we call "low volatility" countries with mean of exchange rate volatility less than 10%, and "high volatility" sub-sample having, $\sigma_{\Delta e}$, more than

10%. As can be seen in table 7, we count 17 low volatility countries and 10 countries with a highly volatility of exchange rate.

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 volatile). Results from IPS tests (reported in Appendix B1, B2 and B3) 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 B5). Almost all of tests lead us to reject the null of non-cointegration.

Now, we can turn to analyzing the long-run pass- through estimates, as reported in Table 8. We begin with the inflation rate as a determinant of ERPT. In view of results, "low inflation" countries experience long run import prices elasticity to exchange rate movement 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[12]. 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)). Also, this result confirms that found in the previous section from the individual estimates (section 5). So, we can conclude that inflation regime is an important source of long-run pass-through heterogeneity across countries.

Next we move to the 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. However, our results show a negative link, with 0.64% long run ERPT in the "less open" economies, which is higher than in the "more open" ones (0.35% by FM-OLS). As we mentioned above, there is no conclusive empirical results in the literature about this phenomenon. For 9 developed countries, McCarthy (2007) shows that association is not particularly strong for ERPT to consumer prices, but there is no evidence that countries with lager import share have a greater ERPT to import prices. However, using panel cointegration approach, Barhoumi (2006) find a positive correlation of pass-through-openness. The main difference with our analysis is the measure of openness used in Barhoumi (2006) which is the tariffs barriers. According to this study, lower tariff barriers countries experience a higher long run pass-through than higher tariff barriers.

[12] Result remains robust when using DOLS method.

Table 8: **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**	0,64**	0,35**	0,47**	0,79**
	[0,49 | 0,57]	[0,70 | 0,81]	[0,60 | 0,68]	[0,27 | 0,44]	[0,43 | 0,52]	[0,74 | 0,84]
DOLS	0,51**	0,82**	0,64**	0,43**	0,39**	0,74**
	[0,46 | 0,55]	[0,76 | 0,89]	[0,60 | 0,67]	[0,34 | 0,53]	[0,35 | 0,43]	[0,69 | 0,79]

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

It is worthwhile to note that Romer (1993) provide an indirect channel, whereby openness is negatively correlated with inflation of consumer prices. In this study, he explains that real depreciation caused by unexpected expansionary monetary policy might be harmful in more open economies (with absence of binding precommitment), thus the benefits of expansion is negatively correlated with the degree of openness. Therefore, monetary authorities in more open countries would expand less which results in lower inflation rates. Nevertheless, our empirical analysis is concerned with pass-through to import prices and not to consumer prices. The main explanation of the negative effect of openness on import prices ERPT is that the greater openness of a country is an indicative of increased competitive pressures between foreign and local producers. In this case, foreign firms are willing to accept adjustments to their markup in order to maintain market share, and the extent of pass-through to import prices would be lower.

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 volatile" countries but import prices increase by 0.74% following one percent nominal depreciation in "high volatile" 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) argues that that pass-through should be less in countries where the exchange rate has been more volatile. According to him, 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 home currency.

7 Has ERPT declined in the Euro Area?

In the final part of the paper, it is useful to focus on the case of the European Monetary Union (EMU) by taking the following sub-sample of 12 countries: Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands and Portugal. It is important to emphasize that the formation of the euro area is likely to have an important impact on ERPT. This is true since the launch of the EMU is seen as a shift in both competition conditions and monetary policy regime.

Empirically, little is said about the European long run ERPT to import prices in context of panel cointegration analysis. As summarized in Table 1, Holmes (2006) examine the pass-through question for 12 European Union members, and his sample involves countries not belonging to the monetary union such as Denmark, UK and Sweden. Also his analysis is concerned with ERPT to consumer prices, and not with the first stage of pass-through, i.e. ERPT to import prices. For 11 euro area countries, De Bandt *et al.* (2007) deal with the micro level of pass-through rather than focusing on aggregate prices reactions, by considering the 1-digit SITC import prices sectors. Therefore, the aim of this section is to combine these two studies, by investigating the degree to which exchange rate variations are transmitted into import prices on the long run and at the aggregate level for 12 countries of the EMU.

It is commonly agreed that the observed decline in pass-through has coincided with general reduced and stable inflation rates in many countries. And consequently, more credible monetary policy regime is seen as a main determinant factor that insulating prices volatility to home currency depreciation. Since the end of Bretton-woods era, European countries have experienced various macroeconomic developments notably in terms of monetary policy and exchange rate regime. This is was started with the "snake in the tunnel" period, followed by the entering to the ERM within the EMS, which has led to the launch of the EMU and the adoption of the euro in January 1999. During this long period of time, it is naturally to see that European countries has gone through different inflation regimes which confirmed by Table 9.

There has been a steady decline in the mean inflation in our 12 euro area countries, which has fallen from 11.4% during the European Snake period to 2.4% over the last decade. It is expectable that this behavior in inflation rate has entailed a decline in the degree of pass-through. With referring to Taylor (2000), ERPT tend to decline in countries where monetary policy shifted strongly towards stabilizing inflation. Thereby, we try to addressee this question, and investigate whether or not pass-through has been diminished since the demise of Bretton-Woods.

Table 9: **Inflation Rates in the EMU (from 1972:2 to 2010:1)**

Country	European Snake 1972:2-1979:1	EMS 1979:2-1990:2	1st and 2nd Stage of EMU 1990:3-1998:4	3rd stage of EMU 1999:1-2010:1
Austria	6,8	3,8	2,6	1,8
Belguim	8,3	4,8	2,2	2
Germany	5,1	3	2,8	1,5
Spain	16,6	10,4	4,6	3,3
Finland	12,3	7,2	2	1,6
France	9,7	7,4	2	1,6
Greece	14,5	19,5	11,7	3,1
Ireland	13,9	9,4	2,3	2,9
Italy	14,4	11,3	4,3	2,2
Luxembourg	7,3	4,7	2,2	2,2
Netherlands	7,7	2,9	2,5	2,1
Portugal	20,6	17,9	6	2,5
Average	**11,4**	**8,5**	**3,8**	**2,2**

Source: OCDE

Now we take a long time series quarterly data from 1973:2 to 2010:1, and then proceed by splitting this sample period into four sub-periods defined according to exchange rate and monetary policy arrangements: First period corresponds to the snake in the tunnel phase from 1972:2 to 1979:1; the second is the SME period from 1979:2 to 1990:2; third period record the launch of the first stage of the EMU and involves also the second stage which finish in 1998:4 ; and the last period corresponds to the formation of the euro area in 1999:1 and lasts until 2010:1. This empirical analysis consist of estimating equation (9) for each of these four sub-periods, in order to assess the ERPT development through different regimes where inflation rate have been considerably declined.

For each sub-period, we conduct IPS tests to check the presence of panel unit root in variables series. So according to the result we are unable to reject the null of non-stationarity for most of series in level (Appendix B4). Thereafter, we test for cointegration relationships between the variables for the four sub-periods. As seen in Table 10, *Group PP* and *Group ADF* Pedroni tests reject the null of non-cointegration in favor of the alternative of cointegration for all countries.

Table 10: **Pedroni cointegration tests**

	European Snake 1972:2-1979:1	EMS 1979:2-1990:2	1st and 2nd Stage of EMU 1990:3-1998:4	3rd stage of EMU 1999:1-2010:1
Group pp-stat	-3.268**	-4.203**	-6.093**	-2.271**
Group adf-stat	-3.141**	-4.779**	-4.225**	-3.110**

Note: Group pp and Group ADF 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.

Now we return to the estimation results as shown in Table 11. As expected, both FM-OLS and DOLS estimators give a strong evidence of a decline of long run ERPT throughout sub-sample periods. During the snake in the tunnel period, import prices responsiveness was higher equaling 0.90% following one percent currency depreciation (by FM-OLS), and this is the utmost pass-through elasticity recorded among the four sub-periods. It is interesting to note that this highest ERPT coefficient occurs in period where mean of inflation rates exceeding the 10% percent in our sample. The moving to SME arrangement does not change considerable the degree to which exchange rate movements affect import prices. Over 1979:2 - 1990:2, ERPT still a little bit higher and upper to 0.80% referring to DOLS estimate. Inflation rates remain higher during this sub-period more than 8% in average. For the third period (first and second EMU stage), exchange rate depreciation is transmitted in a lesser extent, so import prices increase by only 0.60%. It is worthwhile noting that this lowering in pass-through coincides with a substantial fall in European inflation rates. This result advocates for plausible association between inflation and ERPT. Similarly, since the adoption of the common currency in 1999 pass-through remains lower than European snake and SME periods. We obtain 0.50% (respectively 0.53%) by FM-OLS (by DOLS) as import prices reactions. We can add that pass-through elasticities are not quite different in comparison with 1990-3 - 1998:4 sub-period (Figure 3).

Table 11: **Long run ERPT into imports prices in the EMU**

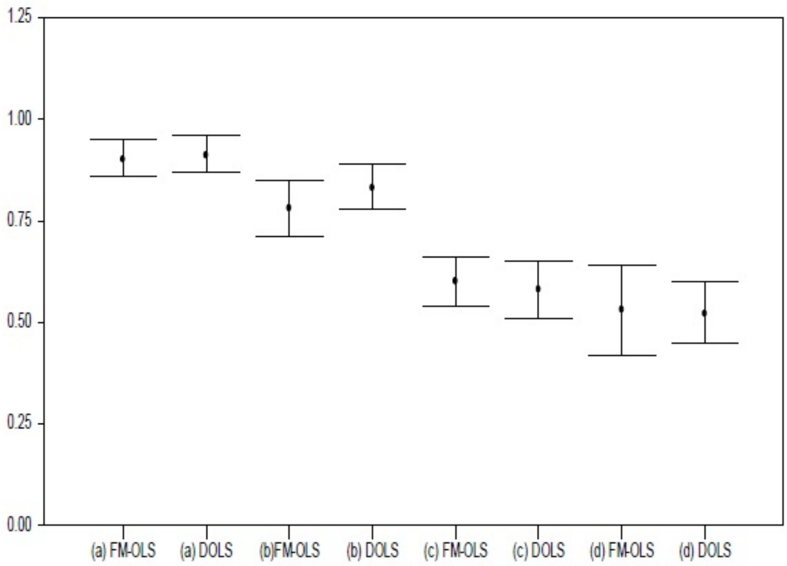
	European Snake 1972:2-1979:1	European Monetary System 1979:2-1990:2	1st and 2nd stage of EMU 1990:3-1998:4	3rd stage EMU 1999:1-2010:1
FM-OLS	0,90**	0,78**	0.60**	0.53**
	[0,86 | 0,95]	[0,71 | 0,85]	[0,54 | 0,66]	[0,42 | 0,64]
DOLS	0,91**	0,83**	0.58**	0.52**
	[0,87 | 0,96]	[0,78 | 0,89]	[0,51 | 0,65]	[0,45 | 0,60]
Inflation	**11,4**	**8,5**	**3,8**	**2,2**

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

Given these results, we notice that the broad decline in long run ERPT is concordant to the steady decline in the mean inflation in our 12 euro area countries. In the light of Taylor hypothesis, it is arguable that this behavior in inflation rate has gave rise a decline in the degree of pass-through. Consequently, a possible positive link between ERPT and inflation can be established in our sub-sample of euro area.

These findings are so convincing since the two macro-determinant, i.e. inflation rates and exchange rate volatility have become more stable throughout the whole sample period (1972:2 to 2010:1). Except few troubling events, European currencies have achieving more stability among each other along the different monetary policy transition which ended with the formation of the euro area in 1999. In the same way, EMU members have gained more credibility through a sustained commitment to maintaining low inflation, and this has been enforced by the explicit primary objective of European Central Bank (ECB), i.e. the price stability. Our findings are in line with the suggestion of Devereux *et al.* (2003) who argue that the euro would become the currency of invoicing (LCP), as a result European imports prices will become more insulated from exchange rate volatility, and thereby ERPT would fall upon the introduction of the euro.

Figure 3: **Long run ERPT Estimates for the Euro Area**



Note : (a) European snake period 1972:2-1979:1 ;
(b) SME period 1979:2-1990: 2 ;
(c) 1st and 2nd stage of EMU 1990:3-1998:4 ;
(d) 3rd stage EMU 1999:1-2010:1.

At the end of our analysis, we want to give further evidence about the significant decline in ERPT for EMU countries. For this purpose, we use a rolling window regression approach in order to check how pass-through has changed over the time. So, we consider individual estimations for our 12 countries after verifying cointegrating relationship in the sense of Engle &

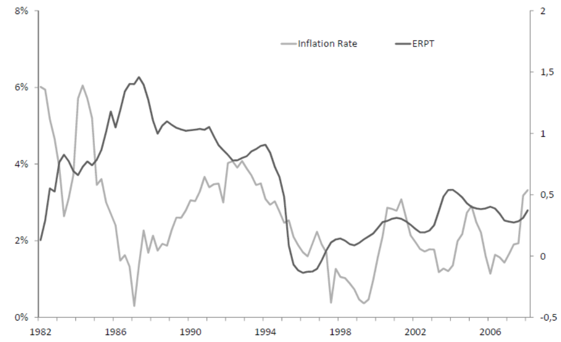
Granger (1987). Long run ERPT estimates are obtained within a fifty-quarter moving window, thereby we estimate equation (9) starting with the window 1970: 1 - 1982:2 and finishing with 1996:2-2008:1. We note that moving window Engle & Granger (1987) tests was conducted for each country (Appendix C2).

Figure 4 display moving window long run pass-through estimates for our 12 euro area countries (moving window estimates with standard error bands are reported in Appendix C1). For all countries except Luxembourg, we note a lower degree of pass-though during the last fifty years in comparison with the eighties. As can be seen, there is a general decline in the extent to which exchange rate fluctuations are transmitted in import prices. It is interesting to note that pass-through has been higher until the EMS crisis in the beginning of the nineties (1992-1993) in which many European currencies have experienced substantial depreciations[13]. Since the launch of second stage of EMU in 1994, there is a strong evidence of lowering ERPT for the most of the euro area members. For example, in Austria import prices reaction was close to the unity during the eighties, while does not exceeding the 0.5% since the formation of the EMU.

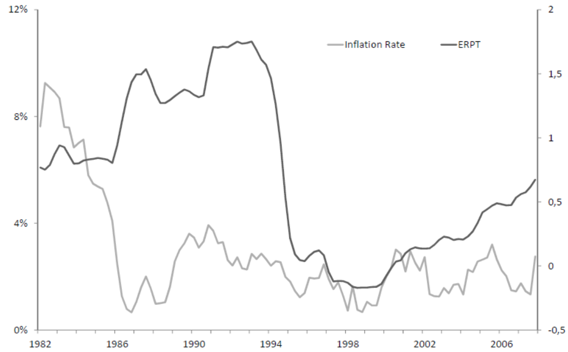
It is known that the European currency has gone through large movements during the first three years of the introduction of the euro, this was resulting in a larger degree pass-through in some of the EMU countries. This rise in ERPT is more evident in Spain, France and Portugal. In Figure 4, we report year to year quarterly inflation rate for each country. We notice that the observed decline in pass-through to import prices was synchronous to the shift towards reduced inflation regime in our sample countries. The visual inspection of Figure 4, show that there is a broad downward tendency for both inflation and ERPT. This is can give a further evidence of the positive correlation between price stability regime and the extent of pass-through. By applying a rolling window technique, Holmes (2008) provides a significant decline in import responsiveness for a panel of 19 African countries. According to his results, this fall in pass-through was concordant to the significant lowering in inflation rate. Consequently a positive association between twenty-year average inflation and ERPT is generally observed.

[13] For example, Italy left ERM in September 1992.

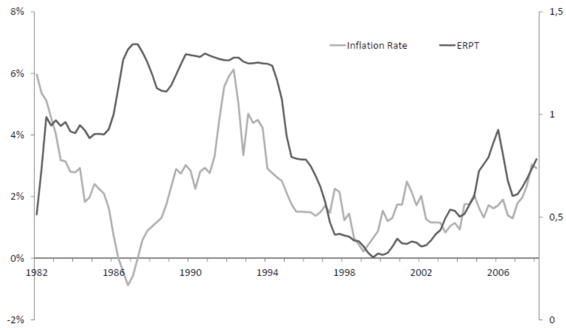
Figure 4: Moving Window long run ERPT and Inflation in the EMU



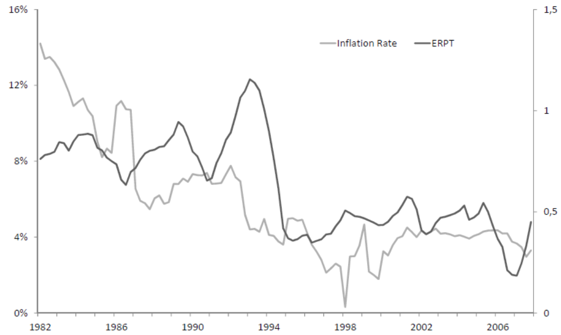
Austria



Belgium

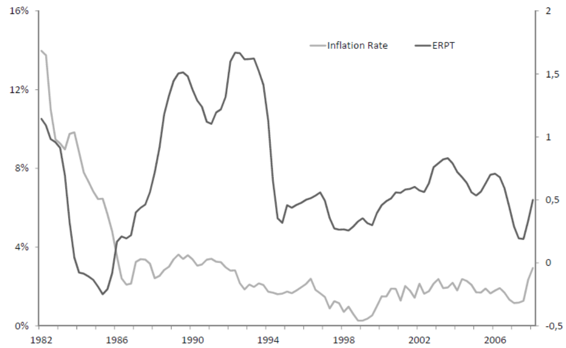


Deutschland



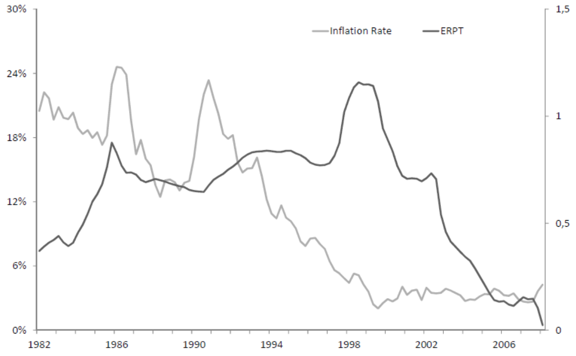
Spain

Finland



France

Figure 4: *Continued*



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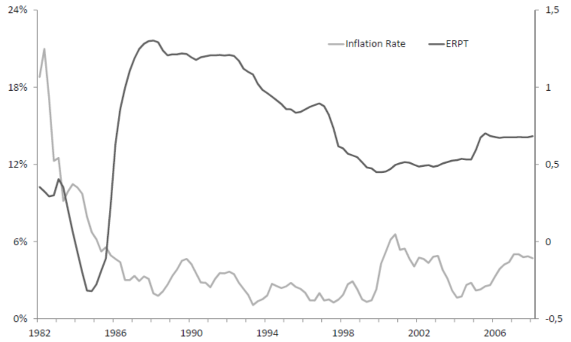

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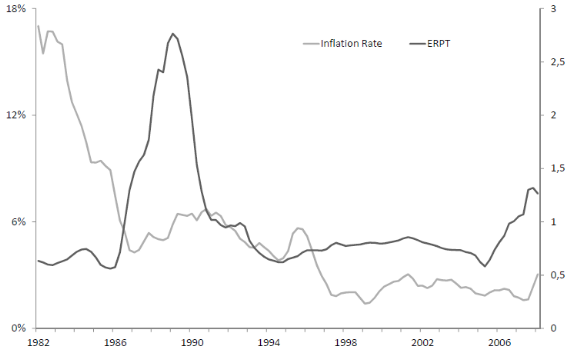

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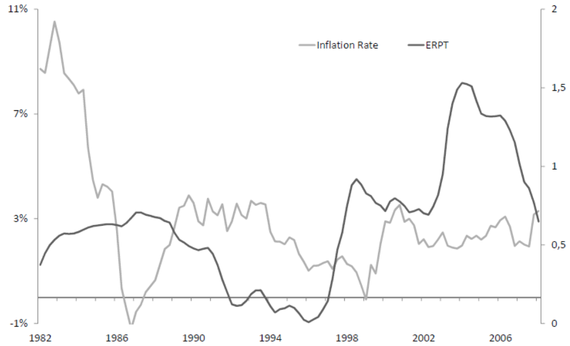
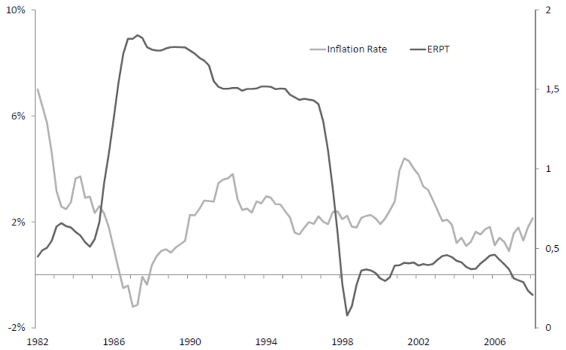
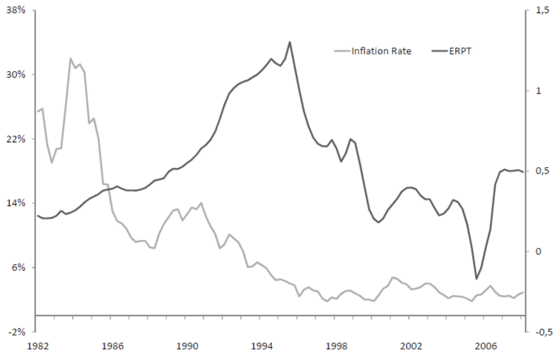


Luxembourg



Netherlands



Portugal

8 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. According to the individual coefficients, there is a significant positive correlation between inflation rates and the extent to which exchange rate variations are passed through import prices.

Then when split our sample in two inflation country regime, we find that "high inflation" countries experienced a higher degree of ERPT than "lower inflation" countries. 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 don't find evidence of positive link between degree of openness and ERPT which is commonly agreed in the pass-through literature.

In the final part of our analysis, we focus on the case of the European Monetary Union (EMU) by taking a sub-sample of 12 euro countries. Our goal is to assess the behavior of ERPT since the collapse of Breton-Woods era and to try to relate it to the change in the inflation environment. As a result, we find a steady decline in the degree of pass-through throughout the different exchange rate arrangements: ERPT elasticity was close to unity during the "snake-in-the tunnel" period while it is about 0.50% since the formation of the euro area. When using a moving window regression for each of 12 countries, a general lowering in the extent of pass-through is confirmed in the EMU. Finally, it is important to emphasize that the observed decline in pass-through to import prices was synchronous to the shift towards reduced inflation regime in our sample of countries. There is a broad downward tendency for both inflation and ERPT. And this is can give a further evidence of the positive correlation between price stability regime and the extent of pass-through.

References

Bailliu, J. & Fujii, E. [2004]. "Exchange Rate Pass-Through and the Inflation Environment in Industrialized Countries: An Empirical Investigation". *Bank of Canada Working Paper No. 2004-21.*

Banerjee, A. & Carrion-i Silvestre, J. [2006]. "Cointegration in Panel Data with Breaks and Cross-Section Dependence". *ECB Working Paper No. 591.*

Barhoumi, K. [2006]. "Differences in Long Run Exchange Rate Pass-Through into Import Prices in Developing Countries: An Empirical Investigation". *Economic Modeling*, 23 (6), pp. 926–951.

Campa, J. & Goldberg, L. [2002]. "Exchange Rate Pass-Through Into Import Prices: A Macro or Micro Phenomenon?" *NBER Working Paper No. 8934.*

Campa, J. & Goldberg, L. [2005]. "Exchange Rate Pass-Through into Import Prices". *The Review of Economics and Statistics*, 87 (4), pp. 679–690.

Choudhri, E. & Hakura, D. [2006]. "Exchange Rate Pass Through to Domestic Prices: Does the Inflationary Environment Matter?" *Journal of International Money and Finance*, 25, pp. 614–639.

De Bandt, O., Banerjee, A. & Kozluk, T. [2007]. "Measuring long run exchange rate pass-through". *Banque de France, Notes d'Etudes et de Recherche No. 173.*

Devereux, M. & Engel, C. [2002]. "Exchange Rate Pass-Through, Exchange Rate Volatility, and Exchange Rate Disconnect". *Journal of Monetary Economics*, 49, pp. 913–940.

Devereux, M., Engel, C. & Tille, C. [2003]. "Exchange Rate Pass-Through and the Welfare Effects of the Euro". *International Economic Review*, 44, pp. 223–242.

Dornbusch, R. [1987]. "Exchange Rates and Prices". *American Economic Review*, 77 (1), pp. 93–106.

Engle, R. & Granger, C. [1987]. "Cointegration and error correction: Representation, estimation, and testing". *Econometrica*, 55, p. 251?27.

Gagnon, J. & Ihrig, J. [2004]. "Monetary Policy and Exchange Rate Pass-Through". *International Journal of Finance and Economics*, 9 (4), pp. 315–38.

Goldberg, P.K. & Knetter, M. [1997]. "Goods Prices and Exchange Rates: What Have We Learned?" *Journal of Economic Literature*, 35, pp. 1243–72.

Hansen, B. E. [1999]. "Threshold effects in non-dynamic panels: Estimation, testing, and inference". *Journal of Econometrics*, 93, pp. 345–368.

Holmes, M. [2006]. "Is A Low-Inflation Environment Associated With Reduced Exchange Rate Pass Through?" *Finnish Economic Papers*, 19 (2), pp. 58–68.

Holmes, M. [2008]. "Long-Run Pass-Through From The Exchange Rate To Import Prices In African Countries". *Journal of Economic Development*, 33 (1), pp. 97–111.

Im, K.S., Pesaran, H., Shin, Y. & Smith, R.J. [2003]. "Testing for unit roots in heterogenous panels". *Journal of Econometrics*, 115, p. 53Ũ74.

Krugman, P. [1987]. "Pricing to Market When the Exchange Rate Changes". In *Real-Financial Linkages Among Open Economies* (S. Arndt and J.D. Richardson. Cambridge, MA.: MIT Press.).

McCarthy, J. [2007]. "Pass-Through of Exchange Rates and Import Prices to Domestic Inflation in Some Industrialized Economies". *Eastern Economic Journal*, 33 (4), pp. 511–537.

Menon, J. [1995]. "Exchange rate pass-through". *Journal of Economic Surveys*, 9, pp. 197–231.

Pedroni, P. [1999]. "Critical values for cointegration tests in heterogenous panels with multiple regressors". *Oxford Bulletin of Economics and Statistics*, 61, p. 653Ũ678.

Pedroni, P. [2001]. "Purchasing power parity tests in cointegrated panels". *Review of Economics and Statistics*, 89, pp. 727–731.

Romer, D. [1993]. "Openness and Inflation: Theory and Evidence". *Quarterly Journal of Economics*, 4, pp. 869–903.

Taylor, J. [2000]. "Low Inflation, Pass-Through and the Pricing Power of Firms". *European Economic Review*, 44, pp. 1389–1408.

Appendix A. Estimation methods

A1. FM-OLS Mean Group Panel Estimator (Pedroni (2001))

We consider the following fixed effect panel cointegrated system:

$$y_{it} = \alpha_i + x_{it}^{'}\beta + \varepsilon_{it}, \qquad t = 1, ...T, \tag{17}$$

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

$$x_{it} = +x_{it-1} + u_{it}, , \forall i \tag{18}$$

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

$$\Omega_{it} = \lim_{T\to\infty} E\left[T^{-1}\left(\sum\nolimits_{t=1}^{T}\xi_{it}\right)\left(\sum\nolimits_{t=1}^{T}\xi_{it}^{'}\right)\right] = \Omega_i^0 + \Gamma_i + \Gamma_i^{'}. \tag{19}$$

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:

$$\hat{\beta}_{GFM} = N^{-1}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)^2\right)^{-1} \times \left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)\,y_{it}^{*} - T\hat{\gamma}_i\right) \tag{20}$$

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 $y_i = \frac{1}{T}\sum\limits_{t=1}^{T} y_{it}$ and $x_i = \frac{1}{T}\sum\limits_{t=1}^{T} x_{it}$ 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:

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

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:

$$t_{\hat{\beta}_{GFM}} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{\beta}_{FM,i}} \tag{22}$$

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

A2. 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:

$$y_{it} = \alpha_i + \beta_i x_{it} + \sum_{k=-K_i}^{K_i} \gamma_{it} \Delta x_{it-k} + \varepsilon_{it}, \tag{23}$$

The group-mean panel DOLS estimator is construct as:

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

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:

$$\hat{\beta}_{D,i} = \left(\sum_{t=1}^{T} Z_{it} Z_{it}' \right)^{-1} \left(\sum_{t=1}^{T} Z_{it} \tilde{y}_i \right) \tag{25}$$

So that the between-dimension estimator can be constructed as

$$\hat{\beta}_{GD} = N^{-1} \sum_{i=1}^{N} \hat{\beta}_{D,i} \tag{26}$$

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

$$\sigma_i^2 = \lim_{T\to\infty} E\left[T^{-1}\left(\sum\nolimits_{t=1}^{T} \varepsilon_{it}\right)^2\right] \tag{27}$$

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

$$t_{\hat{\beta}_{GD}} = N^{-1/2}\sum_{i=1}^{N} t_{\hat{\beta}_{D,i}} \tag{28}$$

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

Appendix B. Stationarity and cointegration tests

B1. Panel unit root tests for different country regime

Table 12: **IPS tests for low inflation countries**

	Low Inflation			
	Level		**First difference**	
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
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

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).

Table 13: **IPS tests for high inflation countries**

	High Inflation			
	Level		**First difference**	
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
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

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).

Table 14: **IPS tests for low openness countries**

	Low Openness			
	Level		**First difference**	
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
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

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).

Table 15: **IPS tests for high openness countries**

	High Openness			
	Level		First difference	
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
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

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).

Table 16: **IPS tests for low volatily exchange rate countries**

	Low Volatility			
	Level		First difference	
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
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

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).

Table 17: **IPS tests for high volatily exchange rate countries**

	High Volatility			
	Level		First difference	
	Intercept	**Intercept and trend**	**Intercept**	**Intercept and trend**
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).

Table 18: **IPS tests for the EMU countries**

	European Snake	**EMS**	**1st and 2nd Stage of EMU**	**3rd stage of EMU**
	1972:2-1979:1	**1979:2-1990:2**	**1990:3-1998:4**	**1999:1-2010:4**
mp_{it}	-1.5022	0.4160	-0.3307	0.0195
e_{it}	0.2934	0.4810	-1.0624	-0.2201
y_{it}	-1.2372	1.5465	-0.7894	0.3976
w^*_{it}	-0.5819	-0.5473	0.0887	-1.0461
Δmp_{it}	-2.450***	-2.579***	-4.1707	-4.8766
Δe_{it}	-2.657***	-2.558	-2.4721	-3.2593
Δy_{it}	-3.310***	-3.1612	-2.7697	-3.0879
Δw^*_{it}	-3.398***	-4.690***	-4.6786	-6.1484

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).

B2. Panel Cointegration Tests

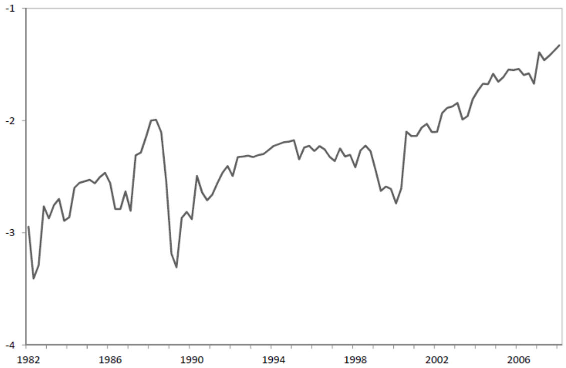
Table 19: **Pedroni tests for different countries regimes**

	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).

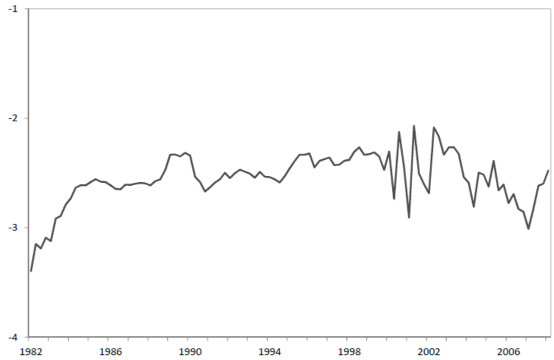
Appendix C. Moving window Engel-Granger Test

Figure 5: Moving window Engel-Granger Test



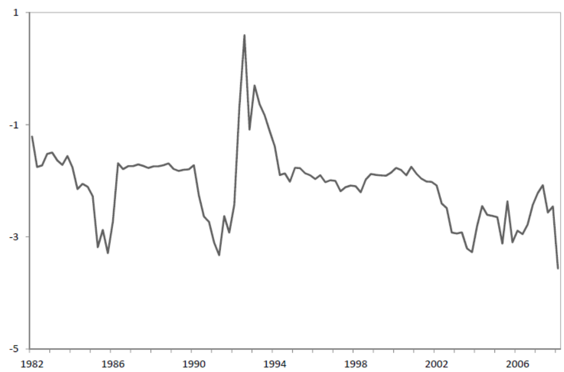
Austria

Belgium

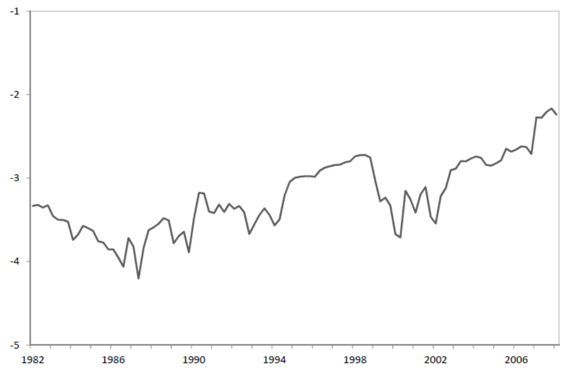


Deutschland

Spain

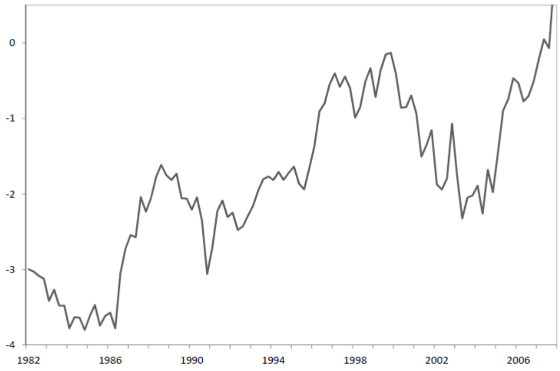


Finland

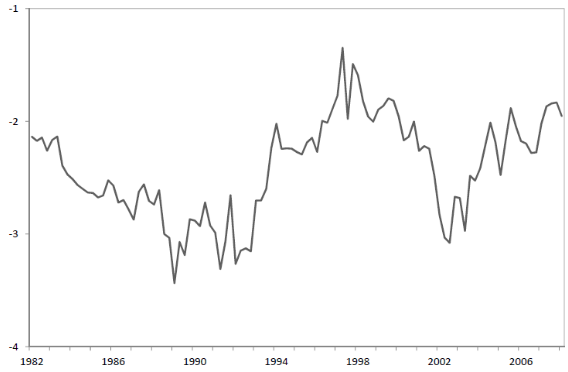


France

Figure 5: *Continued*

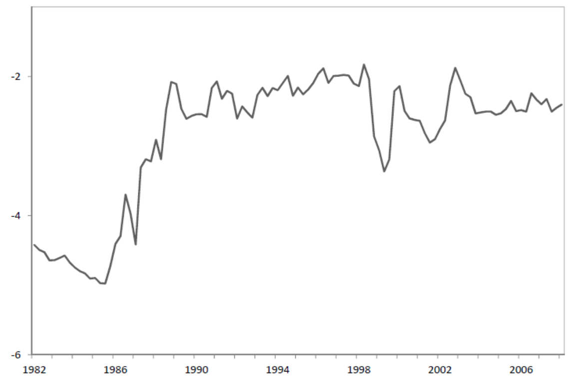


Greece

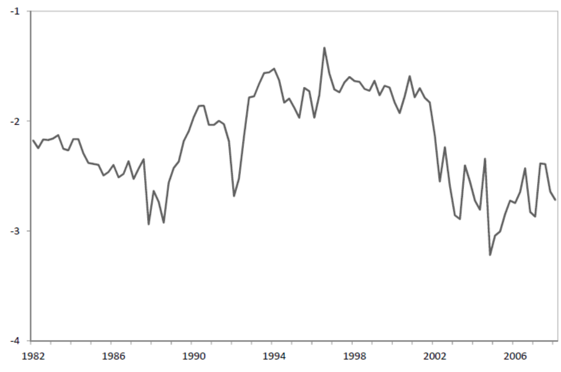


Ireland

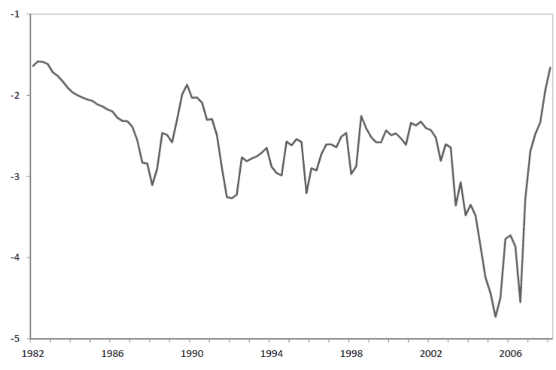
Italy



Luxembourg



Netherlands



Portugal