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Recent Estimates of Exchange Rate Pass-Through to Import Prices in the Euro Area

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Abstract

This paper provides an update on the exchange rate pass-through (ERPT) estimates for 12 Euro area (EA) countries. First, based on quarterly data over the 1990-2012 period, our study does not find a significant heterogeneity in the degree of pass-through across the monetary union members, in contrast to previous empirical studies. As we use a longer time span for the post-EA era than existing studies, this is not surprising, since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions across Euro Area (EA) Member States. Second, when assessing the stability of pass-through elasticities we find very weak evidence of a decline around the inception of the Euro in 1999. However, our results reveal that a downtrend in ERPT estimates became apparent starting from the beginning of the 1990s. This observed decline was synchronous to the shift towards reduced inflation regimes in our sample of countries. Finally, we notice that the distinction between “peripheral” and “core” EA economies in terms of pass-through has significantly decreased over the last two decades.

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

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1. Introduction

The study of the degree of Exchange Rate Pass-Through (ERPT) into import prices is of great policy interest in the Euro area (EA) context. As import prices are a principal channel through which movements in the Euro affect domestic prices and hence also the variability of inflation and output, the issue of pass-through has important implications for divergences in price level developments within the monetary union. A common exchange rate shock may impact EA member states differently depending on their respective relative patterns of external exposure and openness to trade outside the Euro zone. Thus, in achieving its target of medium-term price stability for the whole EA, the single monetary policy of the European Central Bank (ECB) must factor in the extent to which Euro exchange rate changes affect import prices.

It is of special significance that the continuous depreciation of the Euro (about 20 percent on a trade-weighted basis in the first two years) since its introduction has raised concerns that it might increase risks to price stability. The weakening of the exchange rate of the Euro is likely to put upward pressure on import costs and producer prices, which can lead to higher consumer prices. The concern about single currency depreciation affecting price stability has been clearly expressed by the monetary authority in the EA. In fact, the European Central Bank (ECB) cited the inflationary effects of a lower value of the Euro as a factor behind its tightening of monetary policy in 2000.¹ This outcome raised important questions regarding the magnitude and stability of ERPT since 1999, and, mainly, whether EA members will be differentially affected or not by changes in the common external exchange rate. There has been a growing interest in European ERPT in recent years. Studies conducted for the case of EA countries include HÜFNER and SCHRÖDER (2002), HAHN (2003); ANDERTON (2003), CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005), CAMPA and GONZÁLEZ (2006), FARUQEE (2006). A common drawback of these studies is the short time span available since the adoption of the Euro in 1999. Therefore, in our study, we propose an update to ERPT elasticities using a longer time period and more observations for the post-EA era.

Another important issue in the literature is the observed decline in the sensitivity of import prices to exchange rate movements in major industrialized countries. Although the creation of the single currency Euro area constituted a shift in both competition conditions and monetary policy, the European ERPT studies, including CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006), have failed to provide strong evidence of a reduction in pass-through. In fact, there are several factors which may lead to a change in the behaviour of ERPT, and, thus, would explain why the responsiveness of import prices has moved down markedly in the last two decades. An intriguing hypothesis was suggested by TAYLOR (2000), who explains that the shift towards more credible monetary policy and thus, a low-inflation regime would reduce the transmission of exchange rate changes. This assumption is very appealing and has received strong empirical support in the recent literature (see e.g. GAGNON and IHRIG, 2004; BAILLIU and FUJII, 2004; CHOUDHRI and HAKURA, 2006). Nevertheless, the causes of the decline in pass-through are difficult to pin down with certainty, and there is an ongoing debate in this regard. In their sample of 23 OECD countries, CAMPA and GOLDBERG (2005) distinguish “micro-economic” from “macro-economic” explanations. The authors suggested that the product

¹ See the statements given by the ECB in connection to Council monetary policy decisions between February and July 2000.

composition of a country's imports is more important by far than macroeconomic factors such as inflation environment. That is, the shift in the composition of imports towards goods whose prices are less sensitive to exchange rate movements, such as differentiated manufactured products, is the most important driver of the marked fall in pass-through. Given the variability of the empirical findings, we seek here to shed light on some of these issues by revisiting the Euro zone case.

The goal of this paper is to provide new up-to-date estimates of ERPT for 12 EA countries using quarterly data over the 1990-2012 period. First, we begin by estimating a benchmark ERPT equation and analyzing the main properties of the pass-through elasticities in our sample. This enables us to compare our results with those of the existing empirical literature on the EA, such as CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005), CAMPA and GONZÁLEZ (2006). These studies used only a few observations for the monetary union period (post-EA era), thus, their results are updated here. Following this individual estimate exercise, we assess the cross-country differences in our EA sample by investigating whether the inflation level and degree of openness of an economy, as potential macro determinants, determine the magnitude of the pass-through. Next, we verify the stability of the sensitivity of import prices to exchange rate movements over time. There are several reasons to believe that the degree of pass-through has changed since the inception of the Euro in 1999. Among these explanations are the reduction of shares of imports exposed to exchange rate fluctuations and the increase of the choice of the Euro as a currency of denomination. Finally, we estimate our pass-through equation over different time periods and compare results with those obtained over the benchmark period.

The rest of the paper is structured as follows: Section 2 briefly reviews the literature on ERPT. Section 3 provides some theoretical considerations. Section 4 explains the empirical strategy and data sets used. Section 5 reports estimates of exchange rate pass-through to import prices from a benchmark specification, and also investigates for the potential decline in pass-through elasticities. Section 6 offers some concluding remarks.

2. Overview of the literature

The mechanism of ERPT has long been of interest and has spawned many studies over the years. Acknowledging the significant economic literature, we survey only a few important studies concerning pass-through to import prices that are frequently cited. The early literature was mostly composed of papers dealing with ERPT into import prices from a microeconomic perspective.² In that vein, industrial organization characteristics such as the presence of imperfect competition and price discrimination in international markets are the main factors explaining incomplete pass-through. In seminal papers, DORNBUSCH (1987) and KRUGMAN (1987) justified incomplete pass-through as arising from firms that operate in a market characterized by imperfect competition, and adjust their markup in response to an exchange rate shock. As it is well-known, the markup depends on the elasticity of demand for a given product, which is, in turn, determined by competitor prices. Facing a change in the exchange rate, producers can decide whether and to what degree the markup should absorb these changes. When the currency of the importing country is depreciating, a foreign firm might cut its price by reducing its markup, in order to stabilize its price in terms of the

² It is noteworthy that most of the early pass-through literature focused on traded goods prices such as import or export prices and very few on consumer price ERPT.

importing country's currency, in which case pass-through is less than complete. It is important to note that the micro-based literature has a partial-equilibrium approach, *i.e.* it focuses on the response of prices to an exogenous movement in the nominal exchange rate.³

Although the degree of pass-through has played a central role in debates in international economics for a long time, the question of whether pass-through can be influenced by the macroeconomic environment and in particular the role of monetary policy, is a more recent occurrence. The emerging macro literature has focused on the issue of the relatively widespread and on-going decline in ERPT. A popular view in this regard has been put forward in particular by TAYLOR (2000) who provides a model where lower pass-through is caused by lower perceived persistence of inflation. The more persistent inflation is, the less exchange rate movements are perceived to be transitory and the more firms might respond via price adjustments. Thus, countries with credible and anti-inflationary monetary policies tend to experience lower ERPT.⁴ Several empirical studies were very supportive of Taylor's view.⁵ For instance, GAGNON and IHRIG (2004) explore the relationship between pass-through to consumer prices and inflation stabilization in a sample of 20 industrialized countries over the period of 1972-2000. They find that pass-through generally declined in the 1990s and that countries with low and stable inflation rates tend to have low estimated rates of pass-through.

Furthermore, Taylor's hypothesis has been theoretically examined in the context of the new open-economy macroeconomics.⁶ In this type of framework, ERPT will depend on different pricing strategies, *i.e.* whether the foreign exporter follows a producer currency pricing (PCP) or local currency pricing (LCP) strategy. When prices are determined in the exporter's currency (PCP), pass-through tends to be much greater than when prices are set in the importer's currency (LCP). In the extreme case of a purely exogenous exchange rate shock, exchange rate pass-through would be one under producer currency pricing and zero under local currency pricing. It is worth noting that this literature connects macroeconomic and microeconomic factors. DEVEREUX, ENGEL, and TILLE (2003) developed a dynamic general equilibrium model linking the extent of pass-through to monetary policy. They conclude that countries with low relative exchange rate variability and relatively stable monetary policies would have their currencies chosen for transaction invoicing. In this case, prices are sticky in the currency of the importing country (local currency pricing, LCP), and pass-through tends to be low. However, exchange rate pass-through would be higher for importing countries with more volatile monetary policy. Prices will be preset in the currency of the exporter, *i.e.* prevalence of producer currency pricing strategy, and then ERPT will tend to be high.⁷ However, IHRIG, MARAZZI, and ROTHENBERG (2006) warn against the local currency pricing hypothesis. As a matter of fact, exporters may choose to invoice in the currency of the destination market to shield the price paid by its clients from exchange rate movements in the medium-term. However, over the long run, in the case of a protracted

³ As an alternative to this approach, structural vector autoregressions (VAR) have become increasingly popular as a method to estimate exchange rate pass-through (see e.g. MCCARTHY, 2007). One reason for using the structural VAR approach is that it takes explicit account of the endogeneity of the exchange rate and allows for the estimation of pass-through to a set of prices, such as import prices, producer prices and consumer prices, simultaneously.

⁴ This explanation seems to bear more on pass-through to consumer prices than on pass-through to import prices.

⁵ Most of these studies consider the pass-through to consumer prices.

⁶ This strand of literature is based mainly on the OBSTFELD and ROGOFF (1995) seminal Redux model incorporating imperfect competition and price inertia into a dynamic general equilibrium open-economy model.

⁷ A similar finding was obtained by DEVEREUX and ENGEL (2002).

appreciation of the exporter's currency, it will have to adjust its local currency price to keep its margins in the black.

However, there is a serious debate on the prevalence of macroeconomic factors vs. microeconomic factors. GOLDBERG and TILLE (2008) provide empirical evidence suggesting that the choice of invoicing currency is influenced more by the product composition of trade than by macroeconomic factors. If trade is largely homogeneous, the role of macroeconomic variability in invoice currency choice is substantially damped. For producers, the most important driver of invoice currency selection will be the need to have their goods priced the same way as other competing producers price their products. The same view was emphasized by CAMPA and GOLDBERG (2005) in their studies of import-price pass-through in 23 OECD countries. According to the authors, macroeconomic variables - levels of inflation, money growth rates or country size - are weakly correlated with changes in pass-through, and are not of first order importance in explaining pass-through evolution within the OECD over the past 25 years. Furthermore, there is substantial evidence that the shift in the composition of imports towards goods whose prices are less sensitive to exchange rate movements has contributed to a fall in pass-through in many countries in the 1990s. Besides, MARAZZI et al. (2005) take a somewhat different view. According to the authors, "the Campa-Goldberg compositional-change hypothesis" may explain some, but certainly not the lion's share of the decline in pass-through in the United States. This phenomenon can only explain about one-third of the decline in pass-through to U.S. import prices. MARAZZI et al. (2005) provide evidence suggesting that China's surging exports to the U.S. may also be partly responsible for the low levels of observed pass-through in the US economy.

A host of other hypotheses have also been put forward as factors causing incomplete or declining ERPT to import prices. MANN (1986) documented that the increased usage of exchange-rate hedges may shield a firm from exchange rate shocks allowing them to avoid passing such shocks to consumers. Although hedging can allow firms to postpone passing through an exchange rate shock, in the long run a sufficiently large and permanent exchange rate shock will have to be passed through to importers. Another argument for incomplete pass-through is articulated by BODNAR, DUMAS, and MARSTON (2002) which is related to cross-border production arrangements. If production takes place in several stages across many countries, then the costs of producing the final good are incurred in several currencies. This can explain incomplete pass-through as long as all of these currencies do not experience a common appreciation against the export destination's currency. Finally, a recent paper by GUST, LEDUC, and VIGFUSSON (2012) suggests that the process of international globalization itself may induce a fall in pass-through. In their model, lower trade costs (interpreted broadly as increased globalization) increase the exporting firm's relative markup, which in turn allows their prices to be less sensitive to exchange rates yielding lower pass-through.

3. Theoretical Framework

The theoretical framework used here follows FEENSTRA (1989) and COUGHLIN and POLLARD (2004). The model is set in the context of a price-discriminating monopolist and it is a partial equilibrium. Let us consider a domestic importing country that imports a differentiated good q^m from a monopolist foreign firm which is facing competition from a good substitute z in the importing country. Assuming that the differentiated product q^m is weakly separable from other goods in the consumer's utility function, import demand of good

q^m can be expressed as follows: $q^m(p^m, p^z, Y^m)$, where p^m denotes the import price of q^m in the domestic currency, p^z is the domestic currency price of z and Y^m is the income or expenditure on all goods in the importing country. At the same time, the foreign exporter firm produces good q^x for sale in its local market with the following local (foreign) demand: $q^x(p^x, Y^x)$ where p^x is the foreign currency price of the good and Y^x is the income or expenditures on all goods in other countries.

In this economy, the good q is produced only in the foreign country and inputs are allowed to come from both domestic and foreign countries. Thus, factor prices in the foreign country, w^* , will depend on the exchange rate, e (number of units of importing country's currency per unit of foreign currency). The foreign firm's cost function is given by $c(Q, w^*(e))$, where Q is the total quantity produced for both domestic and foreign markets ($Q = q^m + q^x$). Costs are assumed to be homogeneous in degree one in factor prices, so they can be written as $c(Q, w^*(e)) = w^*(e)\phi(Q)$. The foreign firm maximizes profits in its own currency, treating z and Y^m as exogenous.⁸ Then, the profit maximization problem can be stated as:

$$\max_{p^x, p^m} \Pi = p^x q^x + e^{-1} p^m q^m - w^*(e)\phi(Q) \quad (1)$$

The first-order condition for (1) is:

$$p^x = q^x + p^x \frac{\delta q^x}{\delta p^x} - w^* \phi' \frac{\delta q^x}{\delta p^x} = 0 \quad (2)$$

$$p^m = e^{-1} q^m + e^{-1} p^m \frac{\delta q^m}{\delta p^m} - w^* \phi' \frac{\delta q^m}{\delta p^m} = 0 \quad (3)$$

Equations (2) and (3) can be rewritten as:

$$p^x: \frac{\delta q^x}{\delta p^x} \left[p^x \left(1 - \frac{1}{\varepsilon^x} \right) \right] - w^* \phi' = 0 \quad (4)$$

$$p^m: \frac{\delta q^m}{\delta p^m} \left[e^{-1} p^m \left(1 - \frac{1}{\varepsilon^m} \right) \right] - w^* \phi' = 0 \quad (5)$$

where $\varepsilon^i = - \left(\frac{\delta q^i p^i}{\delta p^i q^i} \right)$ is the elasticity of demand with respect to price for $i = x, m$. Knowing that markup over marginal cost is defined as $\mu^i = \left(\frac{\varepsilon^i}{\varepsilon^i - 1} \right)$, the first order conditions become:

$$p^x: \frac{\delta q^x}{\delta p^x} \left[\frac{p^x}{\mu^x} - w^* \phi' \right] = 0 \quad (6)$$

$$p^m: \frac{\delta q^m}{\delta p^m} \left[\frac{e^{-1} p^m}{\mu^m} - w^* \phi' \right] = 0 \quad (7)$$

Therefore:

⁸ Foreign and domestic firms are assumed to act as Bertrand competitors.

$$p^x = w^* \phi' \cdot \mu^x \quad (8)$$

$$p^m = e \cdot w^* \phi' \cdot \mu^m \quad (9)$$

Solving profit maximization yields the standard condition that the price in each market, *i.e.* foreign and domestic, is determined by a market specific markup, μ^i , over common marginal cost, $w^* \phi'$. Our primary focus is on equation (9). This equation shows that the import price p^m (which is expressed in the importing country's currency) depends on three factors: the bilateral exchange rate between importer and exporter, the marginal cost, and the markup of price over marginal cost. Note that the exporter's marginal cost and markup may change independently of the exchange rate. For instance, a change in the cost of a locally provided input (in the foreign country) can shift the marginal cost. Also, adjustments in markups may occur in response to changes in variables specific to the importing country, namely, demand conditions Y^m and the price of the competing product p^z , so that: $\mu^m = \mu^m(Y^m, p^z)$.

Supposing that marginal costs are constant, $w^* \phi'' = 0$, we can derive ERPT elasticity as follows:⁹

$$ERPT = \frac{\delta p^m}{\delta e} \frac{e}{p^m} = \frac{1 + \eta^{w^*e}}{1 - \eta^{\mu m}} \geq 0. \quad (10)$$

where $\eta^{w^*e} = \frac{\delta w^*}{\delta e} \frac{e}{w^*} \leq 0$ and $\eta^{\mu m} = \frac{\delta \mu^m}{\delta p^m} \frac{p^m}{\mu^m} \leq 0$ are the elasticity of import prices with regard to the exchange rate and the elasticity of the markup with regard to the price in domestic country currency, respectively.

According to (10), pass-through elasticity crucially depends on the behavior of marginal cost and markup. In general, ERPT is positive in the sense that a depreciation in the importing country's currency ($\uparrow e$) increases the import price of a good; while an appreciation of the currency value ($\downarrow e$) raises the price of the imported good.¹⁰ Equation (10) suggests that full pass-through ($ERPT = 1$) is a special case. If marginal cost is not affected by exchange rate fluctuations ($\eta^{w^*e} = 0$), *i.e.* the foreign producer uses only local inputs in the production process, and if markup is constant ($\eta^{\mu m} = 0$), pass-through would be complete. In the case of higher sensibility of marginal costs to exchange rate, *i.e.* when ($\eta^{w^*e} = -1$), ERPT will be equal to zero. Besides, in the case of extreme sensibility of markup to domestic currency import price ($\eta^{\mu m} \rightarrow -\infty$), foreign exporters offset exchange rate changes by adjusting markup, and then ERPT tends to zero.

⁹ The derivations of ERPT elasticity are given in more detail in appendix A.

¹⁰ As explained by COUGHLIN and POLLARD (2004), this can be generalized as long as marginal costs are non-decreasing in output, $\phi'' \geq 0$. However, in the case of decreasing marginal costs ($\phi'' < 0$) and an elasticity of input costs with respect to exchange rate inferior to -1 ($\eta^{w^*e} < -1$), ERPT may be negative.

4. Econometric model and data

Here, we focus on the empirical model used to estimate the degree of pass-through, which stems from the analytical framework presented before. As stated by the import price equation (9), in estimating ERPT it is necessary to isolate the exchange rate effect from other effects, i.e. the exporter's cost shifter, importer's demand conditions, and the price of the domestic competitor. Thus, we can present the arguments of the import price equation (9) through a log-linear regression specification similar to that tested throughout the ERPT literature, *i.e.*

$$p_t^m = \beta_0 + \beta_1 e_t + \beta_2 w_t^* + \beta_3 Z_t + \varepsilon_t, \quad (11)$$

where p_t^m are domestic currency import prices, e_t is the exchange rate, w_t^* denotes variable representing exporter costs, and Z_t is a vector including demand conditions and competitor prices in the importing country among other control variables, and ε_t is white noise. As discussed by CAMPA and GOLDBERG (2002), biased estimates of the pass-through coefficient could arise if foreign costs or proxies for markup are correlated with exchange rates but omitted from the regression. Variants of equation (11) are widely used as empirical specifications in the pass-through literature.¹¹

While the general approach is very similar in pass-through studies, there are a few differences between them regarding the specification and the list of control variables. Our primary concern in this study is the pass-through elasticity which corresponds to the coefficient on the exchange rate β_1 in equation (11). β_1 is expected to be bounded between 0 and 1. Specifically, a one-for-one pass-through to changes in import prices, known as a complete ERPT, is given by $\beta_1 = 1$. In this case, exporters let the domestic currency import prices affected by exchange rate move, whereas, when exporters adjust their markup, a partial or incomplete ERPT occurs and $\beta_1 < 1$.

In our investigation the degree of pass-through into import prices is estimated for 12 EA countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain. We consider the same country sample as in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006). However, for the latter studies, time period estimation covers only up until mid-2004. In our analysis, we provide up-to-date ERPT estimates for the main members of the monetary union. The period of estimation corresponds to the interval that spans from 1990:3 to 2012:4 using quarterly data. This allows us to compare our estimates with existing results for EA countries. For each country, data was collected from the IMF's International Financial Statistics and the OECD's Main Economic Indicators and Economic Outlook.

Concerning our dependent variable, *i.e.* domestic import prices, we use the price of non-commodity imports of goods and services. This represents the import prices of core goods by excluding primary raw commodities because of their marked volatility. For the exchange rate of all the countries surveyed we employed the nominal effective trade weighted series, with an increase meaning a depreciation of the national currency, and a decrease meaning an appreciation. Next, the marginal costs of foreign producers are difficult to measure since they

¹¹ See for instance GOLDBERG and KNETTER (1997) for a survey of this literature.

are not directly observable, and thus need to be proxied. A conventional practice is to use a weighted average

of trade partners' costs as in CAMPA and GOLDBERG (2005) and BAILLIU and FUJII (2004). Following this, the foreign costs of each EA country's major trade partners are derived implicitly from the nominal and real effective exchange rate series as follows: $w_t^* \equiv q_t - e_t + ulc_t$, where ulc_t is the domestic unit labour cost (ULC) and q_t is the ULC-based real effective exchange rate. Given that the nominal and real effective exchange rate series are trade weighted, this proxy provides a measure of trading partner costs, with each partner weighted by its importance in the importing country's trade. As regards the foreign firm's markup, in our benchmark specification, we use the output gap, as the difference between actual and HP-filtered gross domestic product (GDP), to proxy for changes in domestic demand conditions.¹² To check the robustness of the benchmark model, in addition to the output gap, we have included the domestic producer prices ppi_t as a proxy for the competitor prices in the importing country (similar to OLIVEI, 2002; BUSSIÈRE, 2012, among others). Additionally, to check the reliability of the output gap as a suitable proxy for the domestic conditions, the real GDP (as in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ, 2005) can be used instead. Furthermore, as is well-known, changes in the exchange rate also influence import prices indirectly through their effects on commodity prices. To consider such a channel as a robustness test we have included oil prices oil_t (in US dollars) as an additional explanatory variable in the pass-through equation. As explained by IHRIG, MARAZZI, and ROTHENBERG (2006), when it was not possible to find the import prices of core goods that exclude all primary raw commodities, the inclusion of commodity prices indexes, such as oil prices, as independent variables should mitigate some of the noise generated by these volatile components. All the robustness tests with different specifications of ERPT equation are reported in Appendix C.¹³

Another concern in the ERPT equation is related to the fact that foreign costs and the exchange rate would have the same coefficient, *i.e.* $\beta_1 = \beta_2$, as predicted by the theoretical framework in HOOPER and MANN (1989). In practice, this restriction does not necessarily hold, since exchange rates are more variable than costs, and thus, the extent to which they are passed onto prices may differ (see ATHUKOROLA and MENON, 1995, for a discussion). To test for the restriction on whether parameters in the exchange rate and foreign costs are equal or not, Wald tests are subsequently conducted.

Finally, we check for the stationarity of our key variables. Augmented Dickey Fuller (ADF) and ZIVOT and ANDREWS (1992) stationary tests reported in Table B.1 in Appendix B indicate that most of the variables are integrated in order one $I(1)$, except the output gap which is by construction a stationary variable.¹⁴ Given that data are non-stationary we investigate the possibility of cointegration between variables in different levels.¹⁵ To achieve

¹² HP-filter of the GDP series (as an estimate of potential) was constructed using a smoothing parameter of 14.400.

¹³ The additional control variables, *i.e.* producer prices and oil prices, are not considered in our benchmark model in order to avoid multicollinearity issues. For instance, we found that the correlation between the output gap measure and the producer prices is quite high.

¹⁴ The ZIVOT and ANDREWS (1992) test allows for one single break under the alternative hypothesis.

¹⁵ Note that we have also have implemented the efficient unit-root test suggested by Elliott and al. (1996), and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS, 1992) test, recently extended by Carrion-i-Silvestre and Sanso (2006), and we have obtained very similar results, *i.e.* the existence of a unit-root in most of the variables. Besides, unit-root tests applied to variables taken in first differences confirm the stationnarity of all variables.

this, in addition to the ENGLE and GRANGER (1987) test (EG hereafter), we also employ the GREGORY and HANSEN (1996) test (henceforth GH) which allows for structural breaks in the cointegrating vector. As reported in Table B.2 in Appendix B.1 there is weak evidence of possible long-run equilibrium relationships among the variables; indeed, the residuals of the ERPT equation in all levels are non-stationary for most of the countries in our sample.¹⁶ This confirms the existing findings of the literature (see CAMPA and GOLDBERG, 2005; CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ, 2005; CAMPA and GONZÁLEZ, 2006, *inter alia*) but with a longer sample of data. Consequently, in what follows first differences of variables are considered. Besides, since data are not seasonally adjusted, quarterly dummy variables are included to capture possible seasonal effects. Eventually, the import-price inflation equation has the following form:

$$\Delta p_t^m = \beta_0 + \beta_1 \Delta e_t + \beta_2 \Delta w_t^* + \beta_3 gap_t + \text{quarterly dummies} + \varepsilon_t, \quad (12)$$

, and is estimated using Generalised Least Squares (GLS) to take into account possible autocorrelation or heteroscedasticity in the residuals.¹⁷

5. Assessing ERPT into import prices

5.1 Results from the benchmark model

Here, the estimation results of equation (12) over the 1990-2012 period are summarized in Table 1.¹⁸ Overall, the estimation results show that the coefficients of the key variables are statistically significant with expected signs, namely exchange rate depreciation and foreign costs positively affect domestic currency import prices. The exception is the output gap which is found to be positively significant only for 4 out of 12 EA countries.¹⁹ This puzzling result has already been pointed out throughout the ERPT literature (see e.g. BUSSIÈRE, 2012). Turning to the estimated ERPT coefficients, we observe that ERPT elasticities are positively significant in all EA countries and bounded between 0.28% (for Austria) and 0.59% (for Italy). Contrary to previous empirical studies, we do not find a wide heterogeneity in the degree of pass-through across the 12 EA countries (see Figure 1). For instance, a significant degree of variability in ERPT estimates across EA countries was reported in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006). Besides, we find that the average of the exchange rate transmission into the aggregate import prices is equal to 0.43%. In other words, a one percent increase in the rate of depreciation of domestic currency raises import prices by 0.43 percent on average in our EA sample.

¹⁶ We have obtained similar results using the well-known cointegration tests of Johansen (1988, 1991), not reported here to save space.

¹⁷ It must be emphasized that when including producer prices in equation (5), the use of instrumental variable estimators may be more accurate. Indeed, domestic firms compete against the exporting firm taking the level of import prices into account, and producer domestic prices may need to be treated as an endogenous regressor (see BUSSIÈRE, 2012). However the implementation of instrumental variable techniques using lagged domestic product prices as instruments shows that the results are very similar to the GLS estimator. This is also confirmed by the implementation of the Hausman test (1978). That's why in what follows parameters are estimated using GLS and not instrumental variable methods.

¹⁸ Because of data availability, the estimation period is 1990:3-2012:3 for Austria and Ireland, and 1990:3-2012:2 for Greece.

¹⁹ Higher domestic demand would tend to raise import prices.

Our estimates of ERPT are slightly lower in comparison with CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006). In the latter papers, the short-run pass-through elasticities are close to 0.66% on average for 11 EA countries.²⁰ This outcome is not surprising since the mentioned studies used fewer observations than us for the EMU era (CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) until mid-2004 and CAMPA and GONZÁLEZ (2006) until the end of 2001). Since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions, finding a relative low and less dispersed ERPT across EA Member States is expected.

Table 1: GLS estimation results from pass-through equation over 1990:3-2012:4 and Wald tests

	Austria	Belgium	Finland	France	Germany	Greece
Constant	0,028 (0,000)	-0,001 (0,808)	-0,006 (0,126)	-0,004 (0,018)	-0,004 (0,026)	0,009 (0,001)
Δe_t	0,287 (0,000) [0,000]	0,428 (0,000) [0,000]	0,323 (0,000) [0,000]	0,372 (0,000) [0,000]	0,379 (0,000) [0,000]	0,476 (0,000) [0,000]
Δw_t^*	0,428 (0,000)	0,607 (0,000)	0,515 (0,000)	0,624 (0,000)	0,583 (0,000)	0,721 (0,000)
gap_t	-0,014 (0,888)	0,311 (0,003)	0,039 (0,622)	0,061 (0,480)	0,024 (0,468)	0,090 (0,266)
Observations	81	82	82	82	82	78
R²	0,891	0,572	0,320	0,653	0,703	0,607
Wald Test	10,363	19,308	10,338	72,496	35,429	42,168
p-value	(0,002)	(0,000)	(0,002)	(0,000)	(0,000)	(0,000)
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Constant	0,006 (0,114)	0,005 (0,026)	0,011 (0,010)	-0,001 (0,737)	0,002 (0,467)	0,000 (0,885)
Δe_t	0,423 (0,000) [0,000]	0,586 (0,000) [0,000]	0,448 (0,000) [0,000]	0,404 (0,000) [0,000]	0,460 (0,000) [0,000]	0,553 (0,000) [0,000]
Δw_t^*	0,329 (0,002)	0,771 (0,000)	0,656 (0,000)	0,637 (0,000)	0,693 (0,000)	0,664 (0,000)
gap_t	0,145 (0,054)	0,170 (0,048)	-0,088 (0,319)	0,119 (0,032)	0,019 (0,776)	0,100 (0,359)
Observations	81	82	82	82	82	82
R²	0,422	0,795	0,292	0,734	0,649	0,590
Wald Test	2,783	29,763	9,756	94,667	50,018	5,647
p-value	(0,099)	(0,000)	(0,003)	(0,000)	(0,000)	(0,020)

Note: Estimations are based on equation (4). Numbers in parentheses are p -values. For the exchange rate coefficient, p -values in parentheses are based on the null hypothesis of zero ERPT, i.e. $H_0: \beta_1 = 0$, while p -

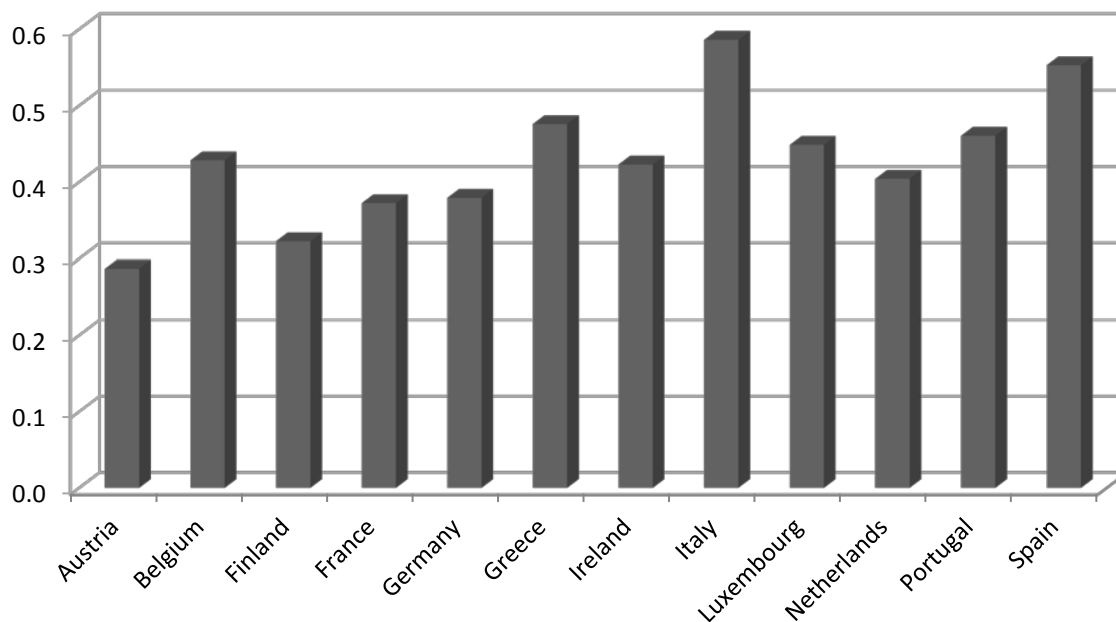
²⁰ In CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006), Belgium and Luxembourg are treated as a single country.

values in square brackets corresponds to the null of full ERPT, i.e. $H_0: \beta_1 = 1$. The Wald test is performed for $H_0: \beta_1 - \beta_2 = 0$.

Moreover, in Table 1, we can test for the prevalence of a local currency pricing (LCP) versus producer currency pricing (PCP) strategy. LCP represents a null hypothesis of zero pass-through, i.e. $H_0: \beta_1 = 0$, whereas PCP implies a pass-through of unity, i.e. $H_0: \beta_1 = 1$. Our results show that both LCP and PCP hypotheses are strongly rejected in all EA countries. According to our results, partial ERPT is the best description for import price responsiveness to exchange rate changes in our country sample. For 23 OECD countries, CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) support this view in the short-run; import price reactions are significantly different from zero in 20 out of 23 countries and significantly different from one for 18 out of 23 countries. However, the authors found that LCP hypothesis is not rejected for Austria, Belgium and Ireland, while the hypothesis of full ERPT (PCP strategy) is accepted for Finland. Nevertheless, the time span in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) covers the period from 1975 through 2003, which contains a longer period prior to the EMU but fewer observations in the post-EA period.

As predicted by some theoretical models we have also tested for the restriction of equality of exchange rate and foreign price coefficients, i.e. $H_0: \beta_1 = \beta_2$ (see HOOPER and MANN, 1989). According to the Wald test results (last 2 lines in Table 1), the hypothesis of equal parameters is rejected for our entire country sample. This outcome is in line with most of the empirical studies which argue that exchange rates are more volatile than costs, and therefore imposing such restrictions does not necessarily hold (see ATHUKOROLA and MENON, 1995).

Figure 1: ERPT elasticities in EA countries over 1990:3-2012:4



Source: Personal calculation.

Regarding the robustness checks, the results obtained from equation (12) seem to be robust to the inclusion of producer prices as an additional explanatory variable to proxy for competitor prices in the importing country (see C.3 in Appendix C). It must be pointed out that the

coefficients of the variable are quite low and not significantly different from zero in all cases except for Greece. Similarly, when we introduce oil prices into the regression, this does not significantly alter the results of the benchmark specification (see Table C.4 in Appendix C). However, we underline that ERPT coefficients are slightly lower when oil prices are introduced into equation (12). This is not surprising, since when commodity prices such as oil prices are excluded from the regression, the pass-through coefficients capture both the direct effect of the exchange rate on import prices and the indirect effect operating through changes in commodity prices. Thus, taking this into account, the latter channel would slightly lower ERPT elasticity (see MARAZZI et al., 2005; IHRIG, MARAZZI, and ROTHENBERG, 2006; MARAZZI and SHEETS, 2007, for a discussion). Finally, we have replaced the output gap by the rate of growth of real GDP in (12) as in CAMPA and GOLDBERG (2005). The results are still the same, *i.e.* the coefficients on output growth are insignificant in most cases, and even when it is the case, it does not affect the other coefficients. We can thus consider that our benchmark specification (12) successfully passed all standard robustness tests.

5.2 Stability of ERPT elasticities

Here, we raise the question of whether ERPT has changed over time in EA countries. Several macro studies have focused on the issue of the widespread and on-going decline in pass-through. This decline has received more attention since it has important implications for the conduct and design of monetary policy. A frequently cited example includes the case of some industrialized countries, namely Canada, Finland, Sweden and the United Kingdom, which experienced a considerable depreciation of the exchange rate in the 1990s without consumer prices being affected as much as was expected. This common experience has led to the widely held belief that pass-through of exchange rate changes into domestic inflation has declined in many of these countries since the 1990s. For our country sample, there are many reasons to expect a change in ERPT behaviour. Especially significant is the fact that the founding of the EA would entail a change in macroeconomic environment and in the competitive conditions (by increasing the share of goods denominated in the single currency), and thus the extent of exchange rate transmission would be affected accordingly. Therefore, it is natural to ask whether the launch of the monetary union in 1999 constitutes a break date in the pass-through mechanism across EA countries.

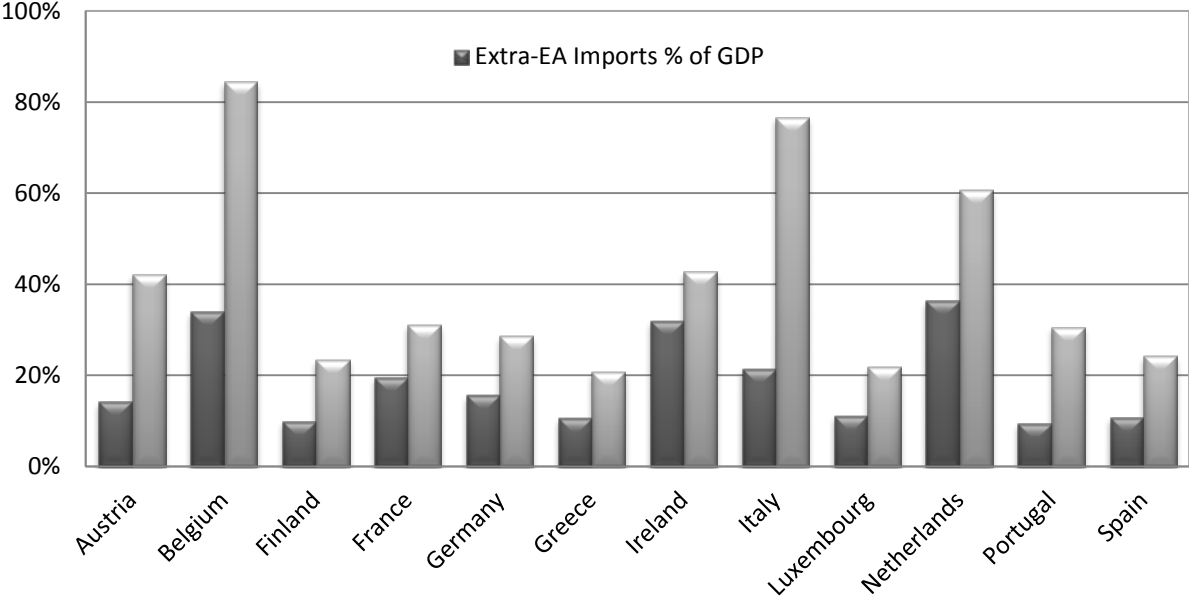
5.2.1 Is there a structural break around 1999?

A number of empirical studies have tested for the presence of a structural break around the date of the inception of the Euro. Using the panel cointegration approach, DE BANDT, BANERJEE, and KOZLUK (2008) provide evidence of a change around the introduction of the common currency (1998-1999) or in the vicinity of the starting of the Euro appreciation against the U.S. dollar (2001-2002). However, CAMPA and GOLDBERG (2005) and CAMPA and GONZÁLEZ (2006) provide weak evidence to back up the existence of a structural break around that time.

In fact, there are some factors that may lead to a change in the rate of ERPT. For instance, the proportion of trade exposed to exchange rate movements diminished after the adoption of the single currency, and this altered the degree of openness in the respective EA countries. For example, as shown in Figure 2, Portugal was more open to trade than Germany over 1990-1998, whereas since the founding of the monetary union it has become less open than Germany. Such developments may lead to a change in the transmission of exchange rate

movements. As explained by DORNBUSCH (1987), pass-through may be higher if the exporters are large in number relative to the presence of local competitors. However the advent of the Euro may have reduced the market power of foreign firms relative to their domestic counterparts, and this may entail a decline in the responsiveness of import prices. Moreover, the choice of the currency of invoicing may have been affected following 1999. Indeed, it is thought that the share of trade being denominated in the Euro has increased. As explained by DEVEREUX, ENGEL, and TILLE (2003), to the extent that the single currency has become the currency of denomination of trade for EA countries, ERPT elasticities may have decreased. To give a further insight into the expansion of the Euro as an invoicing currency across some EA countries, we report in Table 2 the share of imports stemming from outside the EA with prices denominated in Euros. We denote a general increased use of the Euro as the currency of denomination, as it has become a well-established currency (mainly since 2002).

Figure 2: The share of imports in GDP (1999-2012)



Source: Eurostat and the OECD’s Economic Outlook.

For instance, MARAZZI et al. (2005) found that 1997 corresponds to the year after which the decline in U.S. import-price pass-through sped up. Given the large trade flows with Asian countries, the authors argued that the Asian financial crisis of 1997 played a substantial role in the reduction of the pass-through to import prices. They also provide evidence suggesting that the rise in exports from China to the USA may also be partly responsible for the low levels of observed pass-through in recent years.

Therefore, to test for the potential decline in ERPT, as suggested by the above arguments, we performed tests of structural stability in the pass-through rates around the starting of the third stage of EMU, *i.e.* in the vicinity of 1999. To achieve this, we follow CAMPA and GOLDBERG (2005) and CAMPA and GONZÀLEZ (2006) by performing two types of structural change tests on the pass-through coefficients. First, we assume an exogenously imposed break point in 1999 (or close to that date) and perform Chow tests. In a second set of tests we allow for endogenously determined structural break points. Indeed, it is possible that a change in ERPT elasticities did not occur exactly in 1999, therefore Chow tests are also

conducted for a time break around the introduction of the Euro.²¹ Second, to check for the existence of an endogenous break any time over our sample period (1990:1-2012:4), we use ANDREWS (1993) and ANDREWS and PLOBERGER (1994) (AP hereafter) tests without specifying a priori the date at which the change in the ERPT relationship takes place.

Table 2: The share of the Euro as an invoicing currency of EA trade with the rest of the world (%)

Country	Imports of goods										
	2001	2002	2003	2004	2005	2006	2007	2008	2009	2012	2011
Belgium	47,2	53,7	57,8	55,7	51,2	58,3	56,1	56,4	57,7	53	55,7
France	42,6	40,8	44,1	45,7	46,3	44,7	44,8	44,2	44,3	45,5	53
Greece	29,3	35,4	39,2	39,6	32,6	32,3	33,6	37,3	37,9	30,8	33,1
Italy	40,8	44,2	44,5	41,2	39,4	43	44,3	47,8	49,7	46,9	-
Luxembourg	47,2	31,9	41,9	50	43,8	38,8	37,9	38,8	55,3	55	48,7
Portugal	50,3	54,9	58,1	58	54,4	52,6	51,8	53,7	56,6	52,1	45,7
Spain	49,7	55,9	61,1	61,3	56	54,8	56,7	58,8	60,6	59,1	57,7

Source: Review of the international role of the Euro, European Central Bank, July 2012.

The results for the different tests for a structural break are summarized in Table 3.²² Using Chow tests we are not able to reject the null of no structural break for 9 out of 12 EA countries. For these countries, the creation of the monetary union does not affect the extent of pass-through. Only for Belgium, Greece and Ireland is the hypothesis of structural stability rejected, implying that the formation of the Euro area caused a change in exchange rate transmission. Likewise, when applying ANDREWS (1993) and AP tests, we find weak evidence to back up the existence of a (statistically significant) structural break in ERPT into import prices across EA countries. Notable exceptions are Belgium and Italy in the end of 1997, and Greece and Ireland around? 1998. However, these results must be considered with some caution. Indeed, as explained by CAMPA and GONZÁLEZ (2006), the change in ERPT elasticities around 1997-1998 is likely to be related to the negative oil price shock at that time rather than the formation of the Euro zone. Therefore, as a conclusion, it can be emphasized that the presence of a structural break in ERPT coefficients around 1999 does not systematically occur across EA countries.

It is noteworthy that a change in ERPT may not have happened at a specific point in time, such as 1999. Indeed, the decline in exchange rate transmission may be gradual rather than associated with a distinct break date. Furthermore, as discussed by DE BANDT, BANERJEE, and KOZLUK (2008), the changing behavior in the pass-through mechanism may have started before the date of the creation of the Euro (for example during the first or the second stage of the EMU) or after the strengthening of the common currency, in place since 2002. For instance, the acceptance of the Euro as an invoicing currency may be gradual and therefore picked up with a lag as the Euro became well-established. DE BANDT,

²¹ CAMPA and GOLDBERG (2005) and CAMPA and GONZÁLEZ (2006) assume that a structural break might have occurred in May 1998, the month in which the parities among currencies replaced by the Euro were announced.

²² CAMPA and GOLDBERG (2005) and CAMPA and GONZÁLEZ (2006) assume that a structural break may have occurred in May 1998, the month in which the parities among currencies replaced by the Euro were announced.

BANERJEE, and KOZLUK (2008) found that the appreciation of the Euro against the U.S. dollar in 2002 caused a change in the long-run relationship of ERPT.

Table 3: Structural break tests on ERPT elasticities over 1990:3-2012:4

	Austria	Belgium	Germany	Finland	France	Greece
Chow test	0,201	10,183	0,190	1,819	0,062	15,208
	0,904	0,006	0,827	0,162	0,940	0,000
Andrews (1993)	1,366	8,387	2,980	2,651	0,967	8,601
	0,938	0,055	0,558	0,630	0,971	0,050
AP Test	0,222	2,636	0,843	0,640	0,129	2,818
	0,758	0,024	0,249	0,346	0,949	0,019
Break date	-	1997:04	-	-	-	1998:02
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Chow test	5,500	0,741	1,680	0,836	2,459	0,587
	0,064	0,690	0,432	0,658	0,293	0,556
Andrews (1993)	3,898	6,668	4,536	1,974	3,077	2,310
	0,390	0,120	0,301	0,792	0,538	0,710
AP Test	1,177	1,488	0,736	0,280	0,609	0,372
	0,153	0,100	0,295	0,671	0,365	0,559
Break date	1998:03	2007:04	-	-	-	-

Note: Numbers in parenthesis are the p -value of the tests. The test by ANDREWS (1993) uses the maximum of the LM statistics, while the AP test (ANDREWS and PLOBERGE, 1994) uses the geometric mean.

Otherwise, the EMU process entailed some convergence of average inflation rates across the EA members, as a result of efforts to fulfill Maastricht convergence criteria. Thus, the reduction in inflation rates started largely before the inception of the Euro. Given that inflation environment is an important macro determinant of ERPT, one may think that the shift towards more credible and anti-inflationary monetary policy regimes may contribute to lowering the response of import prices to currency movements in the EA. Drawing on this intuition, it is expected that the extent of pass-through was higher in the 1980s than over the course of the last two decades (1990-2012). This might be especially the case of EA countries with historically higher inflation levels, namely Greece, Ireland, Italy, Portugal and Spain. Therefore, we now estimate the ERPT over the 1980s and compare results with those obtained over 1990-2012.

5.2.2 ERPT in the 1980s

A recurrent exercise in the empirical literature is to estimate the ERPT over different sub-sample periods, to test for the conventional wisdom of the decline of pass-through. For instance, the split-sample approach was used by GAGNON and IHRIG (2004) for 20 industrialized countries between 1971 and 2000. The authors estimate the transmission of exchange rate over two sub-sample periods, with break dates chosen based on the observed behaviour of inflation. The first subsample period is a period of high inflation environment, while the second one has lower and more stable inflation. The authors find a strong decline in the pass-through across the two time periods and conclude that this is due to an increased emphasis of the monetary policy on stabilizing inflation. Given the steady decline in inflation

rates in our sample of EA countries, we aim to investigate whether this change in the macroeconomic environment fostered the decline in ERPT. Therefore, we reestimate our benchmark model (12) over 1979:2-1990:2, *i.e.* before the inception of the first stage of the EMU, and compare the pass-through elasticities with those obtained over 1990:3-2012:4.

As reported in Table 4, there are more pronounced cross-differences in ERPT in this period than those recorded over 1990:3-2012:4. Indeed, there were divergent macroeconomic conditions across EA countries during the 1980s, especially between peripheral and core economies. Thus, it is expected that the general process of European convergence, which began before the introduction of the Euro in 1999, would entail a reduction in the variability of pass-through within EA member states. Besides, the hypothesis of null ERPT is rejected for all countries in our sample, while the full ERPT hypothesis is accepted only for Spain. For this country, we observe a higher responsiveness of import prices, *i.e.* when the rate of depreciation increases by 1%, Spanish import price inflation rises by 0.95%. The smallest rate of pass-through is found in Luxembourg, where a one percent rise in exchange rate depreciation leads to an increase in the rate of inflation of import prices by 0.19 percent.

When comparing elasticities estimated in the 1980s (over 1979:2-1990:2) with those in the last two decades (over 1990:3-2012:4), we found a general decline in the rate of pass-through in most EA countries, with the exception of Belgium and Luxembourg (see Figure 3). On average, the import-price pass-through fell from 0.54% over the 1980s to 0.43% over 1990-2012, which corresponds to a decrease of about 0.14% on average.

For their sample of OECD countries, CAMPA and GOLDBERG (2005) compared ERPT estimated over 1975-1989 with the ERPT in 1990-2003 and found that short- and long-run ERPT elasticities declined for 15 out of 21 countries and increased for the other 6 countries. On average, the decline in the short-run import price pass-through is about 0.04 in CAMPA and GOLDBERG (2005), however, the average fall reported in our study is three times as large.

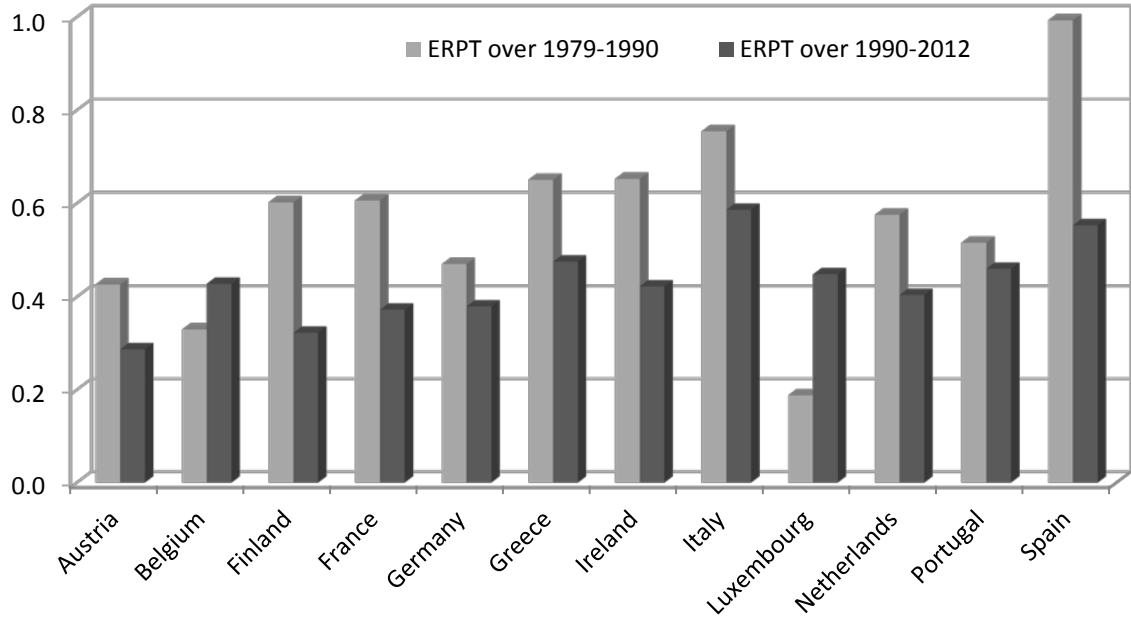
To provide the statistical significance of our results, in Figure 4, we display the point estimates of ERPT with 95% confidence intervals over the two periods. We see that the decline is more pronounced, especially for Spain, Finland and France, where the rates of pass-through are strongly different between the two sample periods. It is interesting to note that Spain had a prior history of high inflation, namely the double-digit inflation rate during the 1970s and the 1980s, whereas in the last two decades the increase in CPI has not exceeded 5% on average. We can posit that this shift towards a stable inflation regime has contributed to the lowering of Spanish pass-through.

Table 4: GLS estimation results over 1979:2-1990:2

	Austria	Belgium	Finland	France	Germany	Greece
Constant	0,007	0,009	0,008	0,006	0,002	0,021
	0,118	0,010	0,108	0,145	0,622	0,006
Δe_t	0,427	0,330	0,602	0,606	0,470	0,650
	0,000	0,000	0,000	0,000	0,000	0,000
	[0,000]	[0,000]	[0,000]	[0,000]	[0,000]	[0,000]
Δw_t^*	0,615	0,309	0,765	0,773	0,661	0,821
	0,000	0,000	0,000	0,000	0,000	0,000
gap_t	0,099	0,570	0,040	0,528	-0,113	-0,051
	0,699	0,000	0,714	0,007	0,509	0,712
Observations	45	45	45	45	45	45
R²	0,712	0,636	0,594	0,772	0,747	0,679
Wald Test	3,567	0,030	5,504	2,935	6,178	1,569
<i>p</i> -value	0,067	0,864	0,024	0,095	0,017	0,218
Chow test	3,419	3,631	5,475	8,510	1,484	8,551
<i>p</i> -value	0,033	0,026	0,004	0,000	0,227	0,000
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Constant	-0,003	0,003	0,010	0,002	0,024	0,027
	0,574	0,685	0,001	0,791	0,001	0,003
Δe_t	0,652	0,755	0,188	0,575	0,515	0,993
	0,000	0,000	0,000	0,000	0,000	0,000
	[0,000]	[0,015]	[0,000]	[0,000]	[0,000]	[0,950]
Δw_t^*	0,884	0,897	0,160	0,774	0,631	1,314
	0,000	0,000	0,064	0,000	0,002	0,000
gap_t	0,109	0,225	-0,067	-0,107	0,271	0,684
	0,571	0,492	0,518	0,684	0,112	0,134
Observations	45	45	45	45	45	45
R²	0,685	0,692	0,494	0,631	0,520	0,747
Wald Test	3,768	0,729	0,189	1,649	0,643	3,332
<i>p</i> -value	0,060	0,398	0,667	0,207	0,428	0,076
Chow test	2,087	2,924	0,918	2,387	13,899	2,732
<i>p</i> -value	0,124	0,054	0,400	0,092	0,000	0,065

Note: Estimations are based on equation (4) over 1979:1-1990:2. Numbers in parentheses are *p*-values. For the exchange rate coefficient, *p*-values in parentheses are based on the null hypothesis of zero ERPT, i.e. $H_0: \beta_1 = 0$, while *p*-values in square brackets corresponds to the null of full ERPT, i.e. $H_0: \beta_1 = 1$. The Wald test is performed for $H_0: \beta_1 - \beta_2 = 0$. The Chow test is performed for the hypothesis that a structural break took place around 1990.

Figure 3: Decline of ERPT into import prices



Source: Personal calculation.

Moreover, it is interesting to note that when we performed Chow tests assuming an exogenously imposed break point around 1990, we found the null of ERPT stability was strongly rejected for most EA countries, except for Ireland and Luxembourg (see last rows in Table 4). These results appear supportive, overall, of the hypothesis of a change in the ERPT mechanism over time. Although the change is not statistically significant for some EA countries, as reported in Figure 4, we can conclude that there has been a tendency towards a decline in pass-through in our sample. The fact that the behavior of pass-through in the last two decades has been different than was the case before seems compelling.

Figure 4: ERPT point estimates

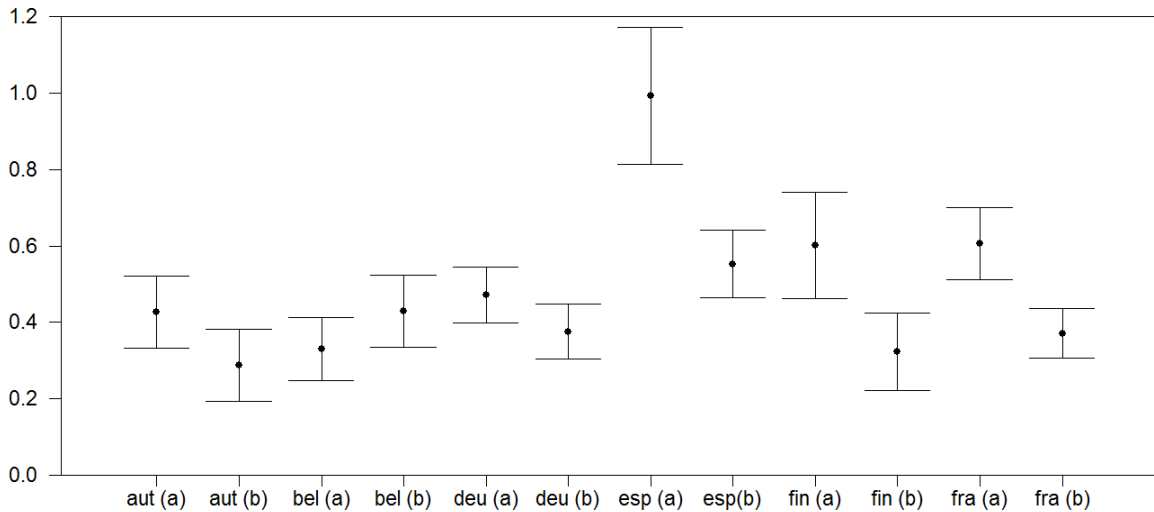
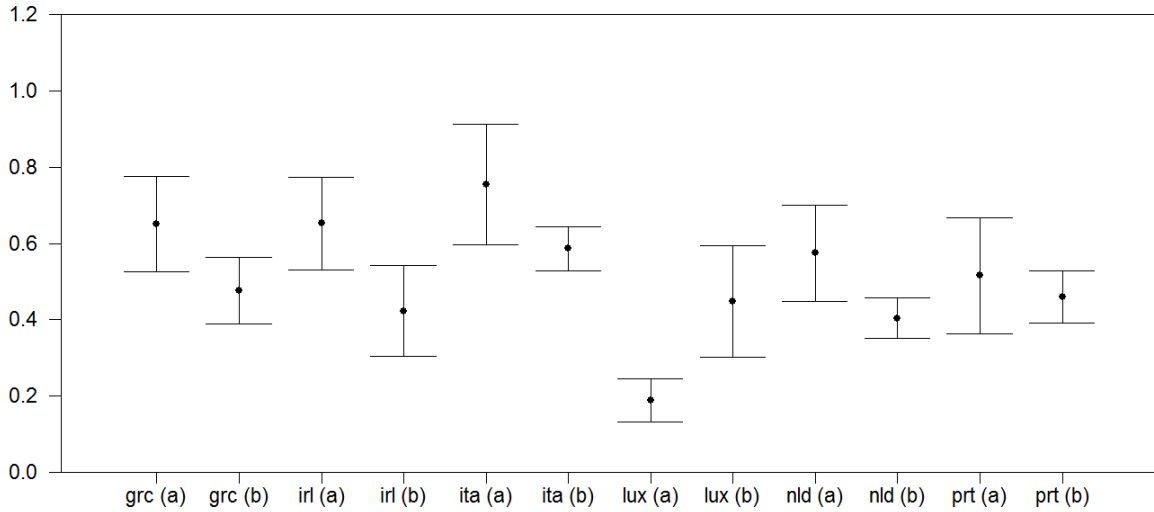


Figure 4: Continued



5.2.3 Evidence from rolling regressions

To give further evidence of the significant decline in ERPT in our sample of countries we use a rolling window regression approach here. This allows us to check how pass-through has changed over time. For this purpose, ERPT elasticities are estimated from equation (4) with a 10-year moving window rolled forward one quarter at a time. We start with the window 1979:2- 1989:1 and finish with 2001:1-2012:4. This will show the evolution of the responsiveness of import prices in EA countries.

The rolling estimates of import price pass-through are shown in Figure 5 (estimates with standard error bands are reported in Figure 6 in Appendix D). We have also reported inflation rates on the same plots to assess whether the shift towards stable inflation environment has been synchronous to the decline in ERPT. For a better understanding of plots in Figure 9, the first observation which lies above 1989:1 (on the horizontal axis) corresponds to the first 10-year sample, *i.e.* the time slot of 1979:2-1989:1.²³ The latest 10-year sample, *i.e.* the 2001:1-2012:4 period, is reported as 2012:4 on the horizontal axis.

A careful inspection of Figure 5 reveals that ERPT to import prices was higher during the 1980s (in the first 10-year window) but appears to trend down afterward in most EA countries, except for Belgium and Luxembourg. The degree of pass-through decreased significantly between our earliest and latest 10-year samples.²⁴ For example, in France, the exchange rate sensitivity of import prices was more than 0.60% in the 1980s, while a steady

²³ The estimates obtained from the first 10-year sample should be close to those displayed in Table 4 in the previous sub-section.

²⁴ As a robustness check, we consider different sample windows, of 15 years for example, in addition to a 10-year window. These robustness tests are important because without them it is not clear whether a change in the pass-through coefficient reflects the new quarters of data entering the sample or the old quarters of data dropping out of the sample. In doing this, we find that the size of the sample window does not really matter. Our results are therefore robust since the decline in ERPT is underlined in most countries of our sample.

decline from 1994 reached 30% of pass-through by the end of 1996. It is interesting to note that pass-through was high until the end of the Exchange Rate Mechanism (ERM) crisis of the European Monetary System in the beginning of the nineties (1992-1993), a time when many European currencies experienced substantial depreciation.²⁵ Since the launch of the second stage of the EMU in 1994, there has been strong evidence of a lowering of ERPT for most EA members. This decline came after the implementation of the Maastricht treaty, which advocated the achievement of a high degree of price stability (among other convergence criteria).²⁶

Our results are in line with IHRIG, MARAZZI, and ROTHENBERG (2006), who report estimates of import price pass-through in G7 countries using a rolling regression framework. For France, they reported that ERPT was about 0.50% and stable through 1996, while in 1997, the estimate began to decrease to under 0.2% by the end of 2004. Among the G7 countries, the authors found the lowest level of import price pass-through at the end of the sample time frame. However, IHRIG, MARAZZI, and ROTHENBERG (2006) explained that this lowering in the rate of pass-through might be correlated to the 1997 Asian financial crisis. Indeed, as discussed in MARAZZI et al. (2005), this explanation would be appropriate for the U.S. case. As a substantial portion of U.S. imports come from Asia, it is expected that the Asian crisis of 1997 may have contributed significantly to the decline in pass-through in the U.S. The authors also provide evidence suggesting that the rising prominence of competition from China may also be partly responsible for the low levels of U.S. ERPT.

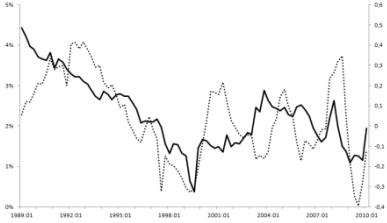
Visual inspection of Figure 5 shows that there is a broad downward tendency for both inflation and ERPT. Finally, we can note that the wide swings of the single currency during the first three years of the monetary union posed a serious threat to price stability in the EA by putting upward pressure on import costs and producer prices.²⁷ Plots in Figure 5 confirm the rise in inflation rates in most EA Members States from 1999 to 2000 due to the extensive depreciation of the Euro. It should be noted that this outcome would explain why ERPT rose in Belgium and Luxembourg instead of decreasing. For the latter countries, the inflation levels were already low, and it is not surprising that pass-through would increase in accordance with the rise of inflationary pressures at the beginning of the EA. Overall, it can be noted that exchange rate changes continue to lead to significant pressures on domestic prices, justifying the growing interest in the issue of pass-through in the context of the EA.

²⁵ For example, Italy left the ERM in September 1992.

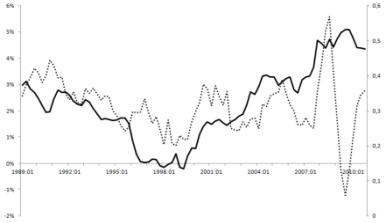
²⁶ As stipulated in the Maastricht convergence criteria, each country's inflation in 1997 had to be less than 1.5 percentage points above the average rate of the three European countries with the lowest inflation over the previous year.

²⁷ See ECB statements by F. Duisenberg (President of the European Central Bank) in 2000.

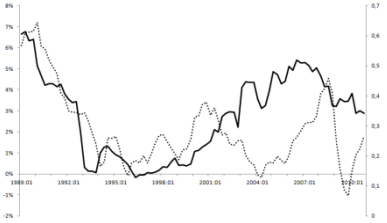
Figure 5: Moving Window ERPT and Inflation in the EA



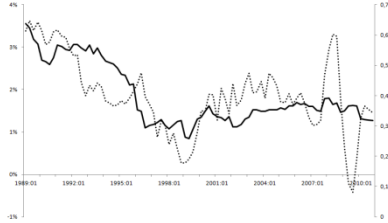
Austria



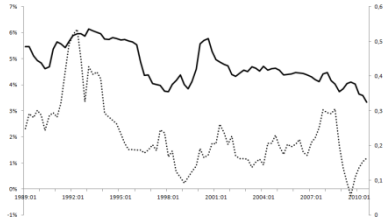
Belgium



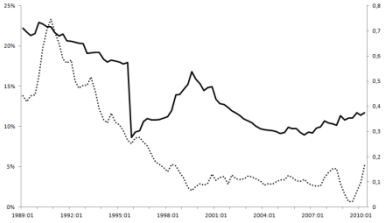
Finland



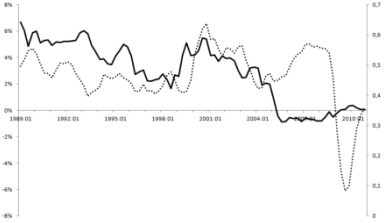
France



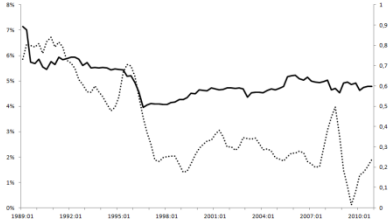
Germany



Greece



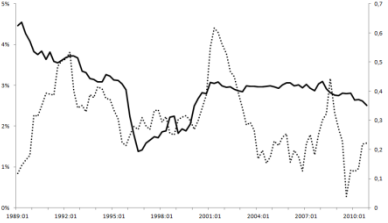
Ireland



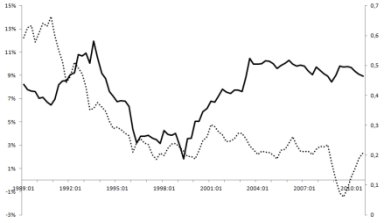
Italy



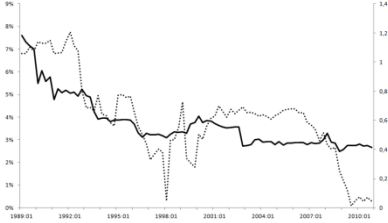
Luxembourg



Netherlands



Portugal



Spain

6. Concluding remarks

In this paper an update of the ERPT estimates is provided for 12 EA countries. First, using quarterly data over the 1990-2012 period, we do not find a wide heterogeneity in the degree of pass-through across the 12 EA countries studied, in contrast to previous empirical works. This is not surprising since previous studies used very few observations for the EA era while in our study the time span for the analysis of the post-EA era is rather long, namely until the end of 2012. In fact, the process of monetary union has entailed some convergence towards more stable macroeconomic conditions, hence finding a relatively low and less dispersed ERPT across EA Member States is expected. Assessing the stability of pass-through elasticities, we find very weak evidence of a decline around 1999.

However, our results reveal that the pass-through estimates appear to trend down from the beginning of the 1990s. We notice that the observed decline was synchronous to the shift towards a reduced inflation regime in our 12 country sample. Besides, when estimating our pass-through equation over 1979:2-1990:2, we obtained more pronounced cross-differences in ERPT than those recorded over 1990:3-2012:4. Indeed, there were divergent macroeconomic conditions across EA countries during the 1980s, especially between “peripheral” and “core” economies. In this way, it is expected that the general process of European convergence, which began before the introduction of the Euro in 1999, would lead to a reduction in the variability of pass-through within EA countries.

Finally, we observe that the wide swings of the single currency during the first three years of the monetary union was a serious threat to price stability in the EA by putting upward pressure on import costs and producer prices. During this period, ERPT increased significantly in some countries, which explains the concerns of the ECB regarding how Euro depreciation affected price stability. Exchange rate changes continue to put significant pressure on domestic prices, justifying the growing interest in the issue of pass-through in the context of the EA.

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Appendix

A. Deriving ERPT elasticity

The profit maximization problem yields the following second order conditions:

$$\begin{vmatrix} \frac{\delta^2 \Pi}{\delta^2 p^x} & \frac{\delta^2 \Pi}{\delta p^x \delta p^m} \\ \frac{\delta^2 \Pi}{\delta p^m \delta p^x} & \frac{\delta^2 \Pi}{\delta^2 p^m} \end{vmatrix} > 0, \frac{\delta^2 \Pi}{\delta^2 p^x} < 0, \text{ and } \frac{\delta^2 \Pi}{\delta^2 p^m} < 0. \quad (13)$$

According to the second inequality in (13):

$$\frac{\delta^2 \Pi}{\delta^2 p^x} = \frac{\delta^2 q^x}{\delta^2 p^x} \left(\frac{p^x}{\mu^x} - w^* \phi' \right) + \frac{\delta q^x}{\delta p^x} \left(\frac{1}{\mu^x} (1 - \eta^{\mu x}) - w^* \phi'' \frac{\delta q^x}{\delta p^x} \right) < 0 \quad (14)$$

where $\eta^{\mu x} = \frac{\delta \mu^x p^x}{\delta p^x \mu^x} \leq 0$, is the elasticity of markup with regard to the price in foreign countries.

By the first order condition (8), $\left[\frac{p^x}{\mu^x} - w^* \phi' \right] = 0$, the sign of (14) depends on the sign of $\frac{\delta q^x}{\delta p^x} \left(\frac{1}{\mu^x} (1 - \eta^{\mu x}) - w^* \phi'' \frac{\delta q^x}{\delta p^x} \right)$. Assuming that demand is well behaved, $\frac{\delta q^x}{\delta p^x} < 0$, we have, $\left(\frac{1}{\mu^x} (1 - \eta^{\mu x}) - w^* \phi'' \frac{\delta q^x}{\delta p^x} \right) > 0$.

Similarly, the third inequality in (13) is expressed as follows:

$$\frac{\delta^2 \Pi}{\delta^2 p^m} = \frac{\delta^2 q^m}{\delta^2 p^m} \left(\frac{e^{-1} p^m}{\mu^m} - w^* \phi' \right) + \frac{\delta q^m}{\delta p^m} \left(\frac{e^{-1}}{\mu^m} (1 - \eta^{\mu m}) - w^* \phi'' \frac{\delta q^m}{\delta p^m} \right) < 0 \quad (15)$$

where $\eta^{\mu m} = \frac{\delta \mu^m p^m}{\delta p^m \mu^m} \leq 0$, is the elasticity of markup with regard to the price in foreign countries.

By the first order condition (9), $\left[\frac{e^{-1} p^m}{\mu^m} - w^* \phi' \right] = 0$, the sign of (15) will depend on the sign of $\frac{\delta q^m}{\delta p^m} \left(\frac{e^{-1}}{\mu^m} (1 - \eta^{\mu m}) - w^* \phi'' \frac{\delta q^m}{\delta p^m} \right)$. If demand is well behaved,, $\frac{\delta q^m}{\delta p^m} < 0$, we have, $\left(\frac{e^{-1}}{\mu^m} (1 - \eta^{\mu m}) - w^* \phi'' \frac{\delta q^m}{\delta p^m} \right) > 0$.

The response of import price p^m with regard to a change in the exchange rate is obtained by using the implicit function theorem to the first-order condition given in the text (equations 8 and 9):

$$\frac{\delta p^m}{\delta e} = \frac{\frac{e \cdot w^* \phi'}{\mu^x} (1 - \eta^{\mu x}) (1 + \eta^{w^* e}) - w^* \phi'' \frac{\delta q^x}{\delta p^x} \frac{p^m}{\mu^m}}{\frac{e(1 - \eta^{\mu x})(1 - \eta^{\mu m})}{\mu^x \mu^m} - w^* \phi'' \left[\frac{(1 - \eta^{\mu x})}{\mu^m} \frac{\delta q^x}{\delta p^x} + \frac{e(1 - \eta^{\mu m})}{\mu^m} \frac{\delta q^x}{\delta p^x} \right]} \quad (16)$$

where $\eta^{w^* e} = \frac{\delta w^*}{\delta e} \frac{e}{w^*}$ is the elasticity of price factors with regard to the exchange rate.

Supposing that marginal costs are constant, $w^* \phi'' = 0$, equation (16) becomes:

$$\frac{\delta p^m}{\delta e} = \frac{e \cdot w^* \phi' (1 + \eta^{w^* e})}{\frac{e}{\mu^m} (1 - \eta^{\mu m})} \quad (17)$$

According to (9):

$$\frac{p^m}{\mu^m} = e \cdot w^* \phi' \quad (18)$$

Therefore:

$$\frac{\delta p^m}{\delta e} = \frac{\frac{p^m}{\mu^m} (1 + \eta^{w^* e})}{\frac{e}{\mu^m} (1 - \eta^{\mu m})} = \frac{p^m (1 + \eta^{w^* e})}{e (1 - \eta^{\mu m})} > 0 \quad (19)$$

Rearranging the latter equation provides ERPT elasticity:

$$ERPT = \frac{\delta p^m}{\delta e} \frac{e}{p^m} = \frac{1 + \eta^{w^* e}}{1 - \eta^{\mu m}} \geq 0. \quad (20)$$

B. Unit-Root and Cointegration Tests

Table A1. ADF and ZIVOT and ANDREWS (1992) Unit Root Tests for main series

	Austria	Belgium	Finland	France	Germany	Greece
Δp_t^m	-3,9454**	-5,6030**	-8,0643**	-6,8245**	-7,1594**	-6,0532**
	-4,85731*	-6,16940**	-8,74938**	-7,09666**	-5,27987*	-7,49465**
Δe_t	-7,0698**	-6,9755	-6,5804**	-6,9895**	-7,0719**	-6,5762**
	-7,61587**	-7,49954**	-6,93960**	-7,50991**	-7,61933**	-7,29540**
Δw_t^*	-6.7092**	-7,0330**	-7,0330**	-7,0029**	-7,0330**	-6.2818
	-7,88917**	-7,88917**	-5.6822	-6.7082	-7.86581**	-4.66986
gap_t	-4,6806**	-4,8614**	-3,4707*	-3,9643**	-4,4618**	-4,1805**
	-3,73447	-5,12743*	-4,49897	-4,22648	-8,17723**	-4,64698
Δppi_t	-6,2127**	-3,40051*	-4,7188**	-4,2540**	-4,11621**	-4,23129**
	-4,69508073	-4,25227	-3,92749	-4,30245*	-4,31624	-3,01229
Δoil_t	-8,3537**	-8,3537	-8,3537**	-8,6739**	-8,3537**	-8,3537**
	-8,94388**	-8,94388**	-8,94388**	-8,94388**	-8,94388**	-8,94388**
Δgdp_t	-4,2438**	-5,1646**	-5,4949**	-4,6927**	-8,7327**	-6,8631**
	-3,58713771	-5,57230**	-6,59565**	-4,08963	-9,27583**	-8,38522**
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Δp_t^m	-8,0565**	-5,5879	-8,4759**	-6,7436**	-5,051**	-5,3442**
	-8,75662**	-5,98038**	-8,95042**	-6,29114**	-6,23993**	-5,62851**
Δe_t	-6,898**	-7,0444**	-6,9755**	-7,0576**	-6,656**	-6,9530**
	-7,41041**	-7,85589**	-7,49954**	-7,61091**	-7,16275**	-7,49144**
Δw_t^*	-7,0330**	-6.4145	-6.6150	-6.9478	-7,0330**	-7.3236
	-7.54180**	-7.989**	-7.29715**	-7.75916**	-5.16994*	-5.04819*
gap_t	-3,0216*	-3,965**	-3,3661*	-3,0151*	-2,645*	-4,0842**
	-4,45079	-4,34674	-4,37522	-3,61456	-4,13153	-4,74697
Δppi_t	-6,8108	-8,3537	-4,560**	-2,960*	-4,43074**	-3,9320**
	-7,29135**	-8,94388**	-4,81311*	-4,21316	-5,16541*	-3,80525
Δoil_t	-3,076*	-5,0258**	-8,3537**	-8,3537**	-8,3537**	-8,3537**
	-4,08098	-5,60308**	-8,94388**	-8,94388**	-8,94388**	-8,94388**
Δgdp_t	-8,3537	-6,492**	-10,4033**	-5,6933**	-6,420**	-3,0679*
	-8,94388**	-7,87149**	-6,23834**	-6,40268**	-4,46148	-4,28633

Note: The first and second row for each series report the ADF and ZIVOT and ANDREWS (1992) test, respectively. **, * denotes rejection of the null hypothesis of the unit root at 5% and 10%, respectively. The ZIVOT and ANDREWS (1992) test allows for one single break under the alternative hypothesis. Lag selection: Akaike (AIC). Maximum lag number = 8.

Table A.2: Engle and Granger (EG), and Gregory and Hansen (GR) Cointegration tests for the benchmark model over 1990-2012

	Austria	Belgium	Finland	France	Germany	Greece
EG Test	-2,18425	-1,44054	-3,50167**	-2,73048	-1,87657	-3,63295**
GH Test						
<i>Break in constant</i>	-4,155	-3,568	-4,423	-4,243	-3,892	-5,147
<i>Break in constant and slope</i>	-5,273	-5,014	-6,132	-6,295*	-6,620*	-5,937
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
EG Test	-3,18447*	-1,57368	-2,78077	-2,65413	-1,65779	-1,9858
GH Test						
<i>Break in constant</i>	-3,728	-5,354*	-3,038	-5,221*	-4,444	-3,864
<i>Break in constant and slope</i>	-3,947	-6,918**	-5,146	-6,887**	-5,601	-5,572

Note: **, * the null hypothesis of the unit root in the residuals (no cointegration) is rejected at 5% and 10%, respectively. The first row reports ENGLE and GRANGER (1987) test. The second row corresponds to GREGORY and HANSEN (1996) tests. Specifications for GH tests include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lag number =8.

C. Robustness checks

B.1 Robustness check with additional explanatory variables

Table B.3: GLS Estimation results with producer prices over 1990-2012

	Austria	Belgium	Finland	France	Germany	Greece
Constant	0,030	-0,007	0,066	-0,001	0,020	0,068
	0,000	0,847	0,494	0,550	0,685	0,000
Δe_t	0,357	0,428	0,418	0,376	0,406	0,423
	0,000	0,000	0,000	0,000	0,000	0,000
Δw_t^*	0,565	0,638	0,651	0,621	0,528	0,699
	0,000	0,000	0,000	0,000	0,000	0,000
gap_t	0,016	0,311	0,063	0,166	0,025	0,133
	0,874	0,003	0,460	0,082	0,457	0,094
ppi_t	-0,233	0,002	-0,022	-0,428	-0,008	-0,020
	0,166	0,859	0,485	0,061	0,633	0,002
Observations	81	82	82	82	82	78
R²	0,898	0,603	0,426	0,653	0,719	0,658
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Constant	0,044	0,002	0,049	-0,003	0,019	-0,002
	0,707	0,955	0,433	0,881	0,393	0,950
Δe_t	0,408	0,590	0,454	0,404	0,462	0,553
	0,000	0,000	0,000	0,000	0,000	0,000
Δw_t^*	0,308	0,783	0,668	0,636	0,699	0,664
	0,004	0,000	0,000	0,000	0,000	0,000
gap_t	0,149	0,172	-0,082	0,118	0,020	0,100
	0,052	0,051	0,357	0,037	0,756	0,364
Δppi_t	-0,012	0,001	-0,013	0,001	-0,006	0,001
	0,746	0,926	0,539	0,904	0,429	0,959
Observations	81	79	82	82	82	82
R²	0,407	0,799	0,296	0,734	0,652	0,590

Note: Estimations are based on equation (12) including the producer prices Δppi_t as an additional explanatory variable. Numbers in parentheses are p -values.

Table B.4: GLS Estimation results with oil prices over 1990-2012

	Austria	Belgium	Finland	France	Germany	Greece
Constant	0,029	-0,001	-0,005	-0,004	-0,003	0,007
	0,000	0,762	0,186	0,042	0,036	0,009
Δe_t	0,244	0,445	0,291	0,303	0,397	0,501
	0,000	0,000	0,000	0,000	0,000	0,000
Δw_t^*	0,321	0,678	0,418	0,451	0,513	0,853
	0,006	0,000	0,002	0,000	0,000	0,000
gap_t	-0,012	0,311	0,031	0,088	0,030	0,056
	0,905	0,003	0,697	0,280	0,362	0,527
oil_t	0,017	-0,006	0,018	0,029	0,012	-0,021
	0,185	0,585	0,289	0,000	0,189	0,086
Observations	81	82	82	82	82	78
R²	0,898	0,604	0,331	0,708	0,711	0,594
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Constant	0,006	0,005	0,011	0,000	0,002	0,000
	0,092	0,019	0,009	0,913	0,341	0,971
Δe_t	0,404	0,578	0,424	0,355	0,432	0,525
	0,000	0,000	0,000	0,000	0,000	0,000
Δw_t^*	0,298	0,745	0,595	0,516	0,613	0,580
	0,029	0,000	0,001	0,000	0,000	0,000
gap_t	0,149	0,179	-0,085	0,118	0,008	0,100
	0,053	0,039	0,342	0,027	0,900	0,358
oil_t	0,002	0,005	0,010	0,020	0,015	0,016
	0,883	0,617	0,610	0,005	0,121	0,232
Observations	81	82	82	82	82	82
R²	0,406	0,794	0,295	0,761	0,661	0,598

Note: Estimations are based on equation (12 including oil prices, Δoil_t , in US dollars as an additional explanatory variable. Numbers in parentheses are p -values.

B.2 Robustness check with alternative proxy for demand conditions

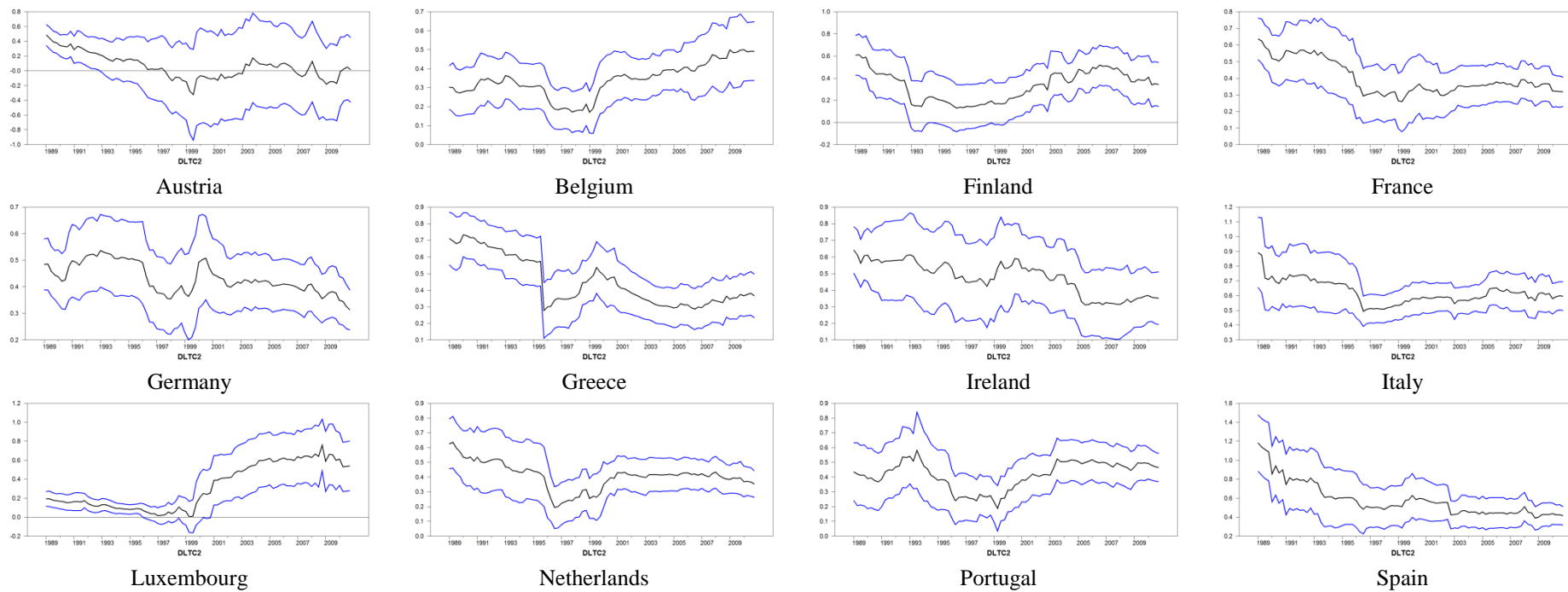
Table B.5: GLS Estimation results with real GDP (growth rate) over 1990-2012

	Austria	Belgium	Finland	France	Germany	Greece
Constant	0,028	-0,003	-0,007	-0,005	0,010	0,009
	0,000	0,290	0,088	0,010	0,838	0,001
Δe_t	0,277	0,399	0,320	0,363	0,411	0,477
	0,000	0,000	0,000	0,000	0,000	0,000
Δw_t^*	0,410	0,550	0,503	0,610	0,481	0,760
	0,000	0,000	0,000	0,000	0,000	0,000
Δgdp_t	0,100	0,572	0,107	0,172	0,138	-0,068
	0,672	0,035	0,470	0,392	0,222	0,324
Observations	81	82	82	82	82	78
R²	0,896	0,539	0,334	0,651	0,715	0,607
	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Constant	0,004	0,004	0,010	-0,001	0,000	-0,001
	0,253	0,052	0,021	0,383	0,900	0,734
Δe_t	0,406	0,568	0,421	0,389	0,449	0,547
	0,000	0,000	0,000	0,000	0,000	0,000
Δw_t^*	0,310	0,745	0,621	0,609	0,673	0,654
	0,005	0,000	0,000	0,000	0,000	0,000
Δgdp_t	0,151	0,253	0,102	0,195	0,304	0,114
	0,165	0,145	0,398	0,113	0,015	0,559
Observations	81	82	82	82	82	82
R²	0,391	0,787	0,289	0,727	0,676	0,594

Note: Estimations are based on equation (12) including the growth rate of real GDP Δgdp_t instead of the output gap as a proxy for the change in the domestic demand. Numbers in parentheses are p -values.

D. Moving window estimates with standard error bands

Figure 6: Moving window ERPT with HAC standard errors



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