Asymmetric responses of prices to exchange rate variations. Evidence from the G7 countries

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Abstract

In this paper, we adopt an asymmetric approach to investigate the response of the CPI to exchange rate changes in the G7 economies during the 1970-2009 period. To do so, we estimate a mark-up model for prices with an asymmetric cointegrating ARDL model, using positive and negative partial sum decompositions of the nominal exchange rates. Our results show that prices react differently to appreciations and depreciations in the long-run, an effect that was previously ignored in the literature, with an important heterogeneity across countries. In particular, in the USA and Germany we find a lower pass-through after an appreciation than after a devaluation, a fact that may be due to downward rigidities. We find the opposite in Canada, France, the United Kingdom and Japan. In Japan this result is consistent with strong entry barriers and suggests that foreign exporters increase pass-through when the Yen is appreciating and decrease it when it is depreciating in order to maintain their market share.

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1 Introduction

Understanding the form and the scale in which nominal exchange rate changes are passed through to domestic prices is an important issue from a monetary policy perspective, with an objective of price stabilisation. In addition, from an international perspective, the exchange rate pass-through (ERPT from now on) determines the level of real exchange rates, an important measure of price competitiveness and a key determinant of the adjustment pattern of the balance of payment.
There is a substantial empirical literature showing that exchange rate changes are less than completely associated with changes in domestic prices at the consumer level (Goldberg and Knetter (1997)). The well established reason is that pricing-to-market by imperfectly competitive firms results in incomplete pass-through to import prices (Dixit and Stiglitz (1977),

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Dornbush (1987), Krugman (1986)). In turn, imported goods play a direct role in consumer prices as a component of the retail price index as well as an inputs of domestic goods. In addition, the extent to which domestic prices are affected by exchange rate fluctuations depends on transport costs, the price stickiness, the importance of non-traded goods in the domestic consumption and the degree of substitution between import and domestic goods.
Now, most empirical research assessing the degree of the ERPT to inflation assumes a symmetric long-run relationship between the price level and the exchange rate. However there are various reasons why this relationship may not be symmetric. For example, an appreciation of the importer's currency may be less passed-through to prices than a depreciation simply because producers are more eager to increase their mark-up than to reduce it. In fact, if downward rigidities exist, then the pass-through measured after an appreciation may be lower than after a devaluation (Bussière (2007)). However if the devaluation takes place in the midst of a recession, then prices might increase less than they would decrease after an appreciation. The reason is that devaluations often results from a downward adjustment of domestic aggregate demand. The resulting recession could act to depress domestic prices, hence implying that domestic prices do not respond much to exchange rate depreciation (Carranza and Gomez-Biscarri (2009), Goldfain and Werlang, 2000).

In sum, well documented downward price rigidities as well as business-cycle factors make the hypothesis of symmetry in the pass-trough unrealistic and too restrictive. Nonetheless, very little research effort has been devoted to asymmetries and non-linearities assumptions of the ERPT. Existing works can be divided in two categories. On the one hand, a few micro-oriented studies focusing on a particular product or industry, use disaggregated data on the industry-level and obtain mixed evidence accross different industries (Yang (2007), Campa and Sebastiá-Barriel (2006), Pollard and Coughlin (2003) and Mahdavi (2002))[1]. On the other hand, the literature exploring the issue from a macroeconomic perspective, is even more scarce (Przystupa and Wróbel (2009) and Nogueira and Leon-Ledesma (2008))[2]. And yet, the asymmetry assumption has crucial macroeconomic implications. If there are asymmetries relative to the direction of exchange rate changes, then the balance of payment does not adjust the same after an appreciation or a depreciation.

Therefore, in this paper, we go further that previous research by adopting an asymmetric approach to investigate the response of the CPI to exchange rate changes in the G7 economies during the 1970-2009 period. The underlying economic theory in which we base our investigation is a mark-up model for prices. From an empirical point of view, we use an asymmetric cointegrating autoregressive distributed lag (ARDL) model that has the enviable advantage to allow long and/ or short-run asymmetry. More precisely, this framework, proposed by

[1] For example, Campa and Sebastiá-Barriel (2006) find that in European manufacturing industries, exchange rate appreciations of the home currency result in a faster adjustment than those caused by a depreciation.

[2] The first paper examines import and consumer prices in Poland while Nogueira and Leon-Ledesma (2008) examine the issue in six countries under an inflation targeting monetary regime

Shin, Yu and Greenwood-Nimmo (2009), consists in a dynamic error correction representation associated with the asymmetric long-run cointegrating regression. Three restricted specifications are obtained by imposing the long and the short-run symmetry restrictions separately or jointly. The resulting empirical model provides elements about the asymmetric response of price level and inflation to exchange rate changes. In total, our model captures long-run effects that were ignored in previous works estimated in first differences only (Bussière (2007), Pollard and Coughlin (2003)).
This paper is organized as follows, Sections 2 and 3 describe the economic theory and the literature dealing with the relationship between prices and the exchange rate as well as the economic justifications for asymmetric pass-through. The methodology and the data used are described in Section 4. Section 5 presents the results regarding the estimation of the symmetric and asymmetric ARDL models growth. Finally, Section 6 provides some concluding remarks.

2 The relationship between prices and exchange rates

This section reviews the theory behind an incomplete pass-through and introduces a linear empirical strategy employed in the literature to assess the degree of ERPT. The restricted assumption of symmetric ERPT in the benchmark model will be relaxed later in the paper.

NIn this sense, it is well established that an incomplete pass-through results from the strategic interaction between firms in an imperfect competition framework (see, for example, Krugman (1986)). Under imperfect competition suppliers have a degree of market power, and set their price taking into account the demand of consumers (Dixit and Stiglitz (1977)). Foreign competitors can thus transfer the exchange rate variation by less than the total magnitude and modify their mark-up according to the domestic demand and their market power. Additional determinants of the degree of exchange rate pass-through include the industrial structure of the economy, the substitutability between domestic and competing imported goods. Intuitively, the less perfect the competition and the more substitutable the domestic and imported goods, the lower the pass-through. On the macroeconomic side, determinants of the degree of ERPT include the trade openness of the economy, the business-cycle, the inflation environment etc (Taylor (2000)). For instance, in a context of recession and high unemployment, the bargaining power of the workers is reduced, implying a lower transfer of an increase in tradable prices to the non tradable sector and resulting in a minor pass-through to the general price index.

A popular empirical model used to estimate ERPT to prices is the mark-up model, derived from the imperfect competition hypothesis[3]. In this model, the dependent variable is the general price index. In the long run, it is a mark-up over total unit costs, including unit

[3] An early empirical application can be found in de Brouwer and Gordon (1998) who use an error-correction specification on Australian data.

labour costs, import prices, and energy prices. Therefore, the explanatory variables include an index of the nominal cost of labour per unit of output, the exchange rate, foreign CPIs or import prices, and an index of primary commodity or oil's prices expressed in dollars. The "pass-through" refers to the direct effect of the exchange rate on the CPI, and is estimated as the coefficient on the exchange rate, β_1 in the empirical model of the following type:

$$p_t = \beta_0 + \beta_1 s_t + \beta_2 p_t^* + \beta_3 oil_t + \beta_3 ulc_t + \varepsilon_t \ , \tag{1}$$

where ε_t is an $i.i.d$ process, p_t is the price index, s_t is the nominal effective exchange rate, p_t^* is the weighted average foreign price index, oil_t is oil's price and ulc is the unit labour cost, all of them in logarithms. According to the mark-up model, in the long run, we should expect all the variables in equation (1) to be positive and significant, implying that the domestic general price level is a mark-up over total unit costs. Foreign inflation is included in the model in order to capture how the international inflationary environment affects the exchange rate pass-through (see Hakura and Choudhri (2001)).

Since the variables in Equation (1) are found non-stationary and cointegrated, the mark-up model can be estimated in an error-correction specification obtained from an autoregressive distributed lag of the CPI on total unit costs as defined above. In addition, some authors (de Brouwer and Gordon (1998), for instance) include the output gap to the short run determinants of the price level, in order to allow the inflation to be modelled as in the Phillips curve framework. de Brouwer and Gordon (1998) observe that an error correction model encompasses a range of economic models for prices and inflation and, at the same time, allows for short-run as well as long-run dynamics. In fact, the long-run determinants of the price level are examined in the long-term relationship (Equation 1) and the inflation rate (the first difference of the CPI) by its short-run determinants (the autoregressive distributed lag of total unit costs plus the output gap). The mark-up model is a used as a reference in modern inflation modelling approach, see de Brouwer and Gordon (1998), Banerjee and Mizen (2007), of England (1999), and Hendry (2001)[4].

3 Economic assumptions justifying asymmetric ERPT

Although exhaustive, a simple linear error correction model obtained from a long-run relationship as the one in Equation (1) assumes a symmetric long-run relationship between the price level and the exchange rate. However there are various reasons why the relationship between prices and exchange rate may be asymmetric. This Section reviews the micro and macroeconomic assumptions justifying an asymmetric relationship.

[4]Other examples of studies employing the mark-up model can be found in Banerjee and Mizen (2007) on US and UK data, as well as by Bruneau, De bandt and Flageollet (2004) and Warmedinger (2004) on Euro area data.

On the microeconomic side, the competitive structure is an influential determinant of the exporters' decisions to pass currency changes into their prices: the higher the firm's market share in the destination country, the lower its incentive to pass nominal appreciation. It is because producers are more eager to increase their mark-up than to reduce it when they can afford it. The most influential factor for the price setter is the price elasticity of demand, which depends on the substitutability between products. Product differentiation appears to relax the competitive pressure of imports, thus leading to monopolistic competition (Bussière (2007)).
In addition, quantity rigidities may imply asymmetric responses. If export quantities are rigid, because they are constrained by trade restrictions or because they work at full capacity, exporters may prefer to keep their price constant in the case of an appreciation of the importer's currency. In fact, the appreciation of the importer's currency reducing the price of exports, increases its demand. In total, the higher quantity rigidities, the lower the incentive to pass nominal appreciation into prices.
A different market structure and the exporter's competitive position may imply an opposite direction of the asymmetry. It is the case if exporters want to maintain their market share after a depreciation that increases their selling price. To do so, they may decide to pass-through a smaller percentage of a depreciation into the import price. Hysteresis effects implying that a temporary loss in market share is likely to become permanent (K. and Klemperer (1989)) amplify asymmetries. A firm that decides to gain competitiveness will therefore adjust its export prices more during appreciations than during depreciations, contrary to what prevails if rigidities hold.
In total, the theoretical literature on pricing-to-market has identified two possible reasons why the elasticity of prices to exchange rate changes may be asymmetric across appreciations and depreciations. If firms are attempting to increase market shares in foreign markets, then more pricing-to-market may occur during appreciations of the exporter's currency. If firms face capacity constraints or are in a favourable competitive position, then the ERPT may be higher during periods of depreciation of the exporters currency (Knetter 1992).

On the macroeconomic side, the position in the business-cycle when the exchange rate variation takes place clearly triggers asymmetric responses. A devaluation often results from a reduction of domestic aggregate demand in the context of a balance of payment adjustment. The resulting recession could act to depress domestic prices, hence implying that a devaluation results in limited inflation.

In sum, the symmetric hypothesis seems to be unrealistic and restricted. Yet, only little research effort has been devoted to include the asymmetry assumption of the ERPT[5]. And most paper have considered short-run dynamic asymmetries only and, as such, abstracted from

[5]see Marston (1990), Swammy and Thruman (1994), Pollard and Coughlin (2003), Campa and Sebastiá-Barriel (2006), Yang (2007), Bussière (2007)

long-run nonlinearity. Yet, this seems to be too restricting since almost all the assumptions just mentioned justify the existence of asymmetries in the long-run also. Indeed, the industrial structure of the economy and the nature of the competition context are long-run structural features.

To our knowledge, only two works have explored long-run asymmetries in an appropriate cointegrated framework. The first one, Webber (1999), explores the question in a cointegrated framework for 8 Asian countries using the Engle and Granger and the Johansen procedures. He finds that the response of import prices is bigger for depreciations than appreciations. Similarly, Wickremasinghe and Silvapulle (2004) examine ERPT on manufactured import prices of Japan using asymmetric unit root and cointegration tests and asymmetric models. They find evidence for cointegration and an asymmetric estimated pass-through coefficient of 98% and 83% for appreciations and depreciations, respectively. In the short-term relationship, only depreciations are significant determinants of import prices during the total sample.

In total, literature dealing with the consumer price hence exploring the issue from a macroeconomic perspective is scarce. On the other hand, most papers consider only short-run dynamic asymmetries and abstract from long-run nonlinearity. To address these issues, our solution is to adopt an asymmetric cointegrating ARDL model that has the great advantage to allow asymmetry in the underlying long-run relationship, in the short-run dynamics or in both. The following Section introduces the specification that we investigate in this paper.

4 Methodology

As it was mentioned before, the majority of the literature which analyzes the ERPT maintains the assumption that the underlying cointegrating relationship may be represented as a linear combination of the underlying nonstationary variables (equation 1).

The standard approach would be to test if there exists a linear cointegration relationship in Equation (1). A very convenient framework to do so is the ARDL framework , suggested by Pesaran, Smith and Shin (2001) and Pesaran, Smith and Shin (1996). The choice of this approach is based on the following considerations. Firstly, unlike most of the conventional multivariate cointegration procedures, which are valid for large sample sizes, the bound test is suitable for a small sample size study. Secondly, the bound test does not impose restrictive assumptions that all the variables under study must be integrated of the same order. Its asymptotic distribution for the F-statistic is non-standard under the null hypothesis of no cointegration relationship between the examined variables, irrespective whether the explanatory variables are purely $I(0)$ or $I(1)$, or mutually cointegrated.

One interesting feature of the ARDL model is that it takes into account the error correction term in its lagged period. The analysis of error corrections and autoregressive lags fully covers both the long-run and short-run relationships of the variables tested. As the error correction term in the ARDL does not have restricted error corrections, the ARDL is called an “Unrestricted Error Correction Model".

Defining the endogenous variable as $\Delta p_t = p_t - p_{t-1}$, $\Delta s_t = s_t - s_{t-1}$ as the rate of appreciation/depreciation, and a row of control variables matrix as $Y_t = (p_t^*, oil_t, ulc_t)'$, in order to test for linear cointegration, we can consider the following linear error correction multi-variance VAR model:

$$\begin{aligned}\Delta p_t =& \mu + \{\rho_p p_{t-1} + \rho_s s_{t-1} + \rho_Y Y_{t-1}\} \\ &+ \sum_{i=1}^{p} \Psi_i' \Delta p_{t-i} + \sum_{i=0}^{p} \phi_i' \Delta s_t + \sum_{i=0}^{p} \theta_i' \Delta Y_t + \sum_{i=0}^{p} \gamma_i' \Delta Z_t \upsilon_t \end{aligned} \tag{2}$$

with μ being an unknown vector of intercept, Z_t a vector of variables which may have effects in the short run but are not are not included in the cointegration analysis above (the output gap, for example), and $\upsilon_t \sim IN(0, \sigma^2)$. The bounds test in Equation (2) can be performed using two separate statistics. The first involves an F-test on the joint null hypothesis that the coefficients on the level variables are jointly equal to zero. The second is a t-test on the lagged level dependent variable.

It is important to notice that instead of the conventional critical values, this test involves two asymptotic critical value bounds, depending on the dimension and cointegration rank of the forcing variables Y_t' and s_t as well as on whether restrictions are placed on the intercept and trend in the model. In particular, Pesaran et al. (2001) show that the critical values take on lower and upper bounds providing then bounds covering all possible classifications of into $I(0)$, $I(1)$ and mutually cointegrated processes. The PSS test is then formulated as follows:

$$\begin{gathered} H_0 : \rho_p = 0, \quad H_0 : \rho_s = 0 \quad \text{and} \quad H_0 : \rho_Y = 0 \\ \text{against} \\ H_0 : \rho_p \neq 0, \quad H_0 : \rho_s \quad \text{and} \quad H_0 : \rho_Y \neq 0 \end{gathered}$$

If the computed statistic exceeds its respective upper critical values, then there is evidence of a long-run relationship regardless of the order of integration of the variables. If, in contrast, it is below, it is not possible to reject the null hypothesis of no cointegration and if it lies between the bounds, any inference is inconclusive. Yet, if the test statistic exceeds its upper bound, then we reject the null of no cointegration. The procedure then amounts to test the assumption of no relationship (in levels) between the dependent variable p and the independent variables s_{t-1} and Y_{t-1}' in the regression.[6]

Nonetheless, as suggested by Shin et al. (2009), the previous symmetric linear combination of nonstationary stochastic regressors might be overly restrictive. Instead of imposing this

[6]Some important issues in the specification are whether to include a deterministic trend in the ECM, and the number of lags, p, to include in the regression. Since the test is based on the assumption that the disturbances are serially uncorrelated, this is not a trivial issue. Thus, as noted by Pesaran et al. (2001), for the bounds test to be valid, it is especially important to ensure that there is no serial correlation.

symmetry, they propose to analyzed the long-run dynamics allowing for asymmetries in the context of nonlinear error correction models. Following Schorderet (2004) and Shin et al. (2009), the starting point consists in the decomposition of a time series s_t into its positive (s_t^+) and negative (s_t^+) partial sums. In our particular case, this corresponds to:

$$\mathbf{s}_t^+ = \sum_{j=1}^{t} \Delta s_t^+ = \sum_{j=1}^{t} \max(\Delta s_t, 0), \quad \mathbf{s}_t^- = \sum_{j=1}^{t} \Delta s_t^- = \sum_{j=1}^{t} \min(\Delta s_t, 0) \ , \tag{3}$$

where Δs_t^+ and Δs_t^- are the partial sum processes of depreciations and appreciations, respectively. Following Shin et al. (2009), Equation (2) can then be modified to allow for asymmetric cointegration, such that the specification becomes:

$$\begin{aligned} \Delta p_t =& c_o + \{\pi_p p_{t-1} + \pi_s^+ s_{t-1}^+ + \pi_s^- s_{t-1}^- + \pi_y Y_{t-1}\} \\ &+ \sum_{i=1}^{p} \Psi_i' \Delta p_{t-i} + \sum_{i=0}^{p} \phi_i^{+'} \Delta s_t^+ + \sum_{i=0}^{p} \phi_i^{-'} \Delta s_t^- + \sum_{i=0}^{p} \theta_i' \Delta Y_t + \sum_{i=0}^{p} \gamma_i' \Delta Z_t \upsilon_t + \upsilon_t \ , \end{aligned} \tag{4}$$

where the variables are defined as in Equation (2), the superscripts '+' and '-' denote positive and negative partial sums computed following the previously mentioned decomposition method. In (4), the relationships is defined such as to allow for cointegration between prices and the positive and negative components of the exchange rate, controlling for the rest of the underlying variables.

Equation (4) opens the possibility that the process being modelled may exhibit asymmetries in both the short- and the long-run, only in the long-run or only in the short-run. Indeed, whereas the first part part in the equation represents the long-run relationship which you can evaluate by bounds testing following Pesaran et al. (2001), the second line contains the lags of the asymmetric exchange rate terms in first differences - it is on this part of the equation that we will focus when testing for short-run asymmetry.

More in detail, as in the linear cointegration case, in Equation (1), we can bound test for a long-run relationship among the variables. Shin et al. (2009) suggest the following operational tests for the existence of an asymmetric (cointegrating) long-run relationship. Based on the error correction model, if $\pi_p = 0$ Eq. (4) reduces to the linear regression involving only first differences, thus implying that there is no long-run relationship between the levels of p_t, s_t^+, s_t^- and Y_t'. This test can be performed either as t-statistic testing $\pi_p = 0$ or as an F-test of the joint null hypothesis $\pi_p = \pi_s^+ = \pi_s^- = \pi_Y = 0$. The asymptotic distributions of these three statistics, denoted t_I and F_I. As before, the critical values tabulated by Pesaran et al. (2001) provide critical value bounds for all classifications, irrespective of whether the regressors are I(0), I(1) or mutually cointegrated.

In addition, symmetry in the short-run can be tested either by a strong form, which implies $\phi_{t-h}^{+'} = \phi_{t-h}^{-'}$ or by a weaker form called the additive short-run symmetry which implies $\sum_{i=1}^{q-1} \phi_{t-i}^{+'} = \sum_{i=1}^{q-1} \phi_{t-i}^{-'}$. In both cases, a standard Wald test, either on each pairwise combination of '+' and '-' dynamic coefficients or the sum of each one of them, can be performed.

Finally, Having estimated the model defined in (4) it is now possible to compute the long-run multipliers as follows: $L_s^+ = \pi_s^+/-\pi_p$, $L_s^- = \pi_s^-/-\pi_p$ and $L_Y = \pi_Y/-\pi_p$. As before, long-run symmetry can be tested by means of a Wald test, with symmetry implying $L_s^+ = L_s^-$.

5 Data and Specification Results

The dependent variable is the consumer price index (CPI), which we obtained from the International Financial Statistics, IMF. Regarding the explanatory variables, the exchange rate corresponds to the nominal effective exchange rate, extracted from the Bank of International Settlements (BIS). The foreign CPI for each country i is calculated as a weighted average of CPIs in j, with weights corresponding to the share of each partner in average values of imports and exports of goods and services over the 2005-2007 period. [7] Finally, the unit labour cost, the price of oil and the output gap were also obtained from the IMF and the OCDE's economic Outlook. Data are quarterly from 1970:1 to 2009:3 for the G7 countries: Canada, France, Germany, Italy, Japan, the United States and the United Kingdom[8].

In order to explore the symmetric and asymmetric long-term relationship between prices and the exchange rate, controlling for a set of variables, we first performed the bounds contegration test in the possible version of Equation 4: (i) allowing for no asymmetry neither in the short nor in the long run, (ii) allowing for long-run asymmetry but imposing short-run symmetry, (iii) allowing for short-run asymmetry but restricting the log-run to have symmetric effects and, (iv) allowing for both, short and long-run asymmetry. Results from the bounds cointegration test and the Wald test for symmetry are presented in table (1) below.

Regarding the bounds cointegration tests, neither linear nor asymmetric cointegration can be rejected in all countries. Indeed, as revealed from the bounds tests, there is clearly a valid long-run relationship between prices and the explanatory variables in all the countries as both, the t_I and F_I test statistics, exceed their respective upper critical values in most of the cases[9]finding that prices are cointegrated with a combination of input prices confirms existing evidence for the G7 countries[10].

Tables (2) to (3) report the estimation results for the ARDL regression for the symmetric and asymmetric models[11]. As seen, there is a pronounced positive association between prices

[7] In the weighting matrix we only consider OECD countries.

[8] Due to data availability, the analysis for Germany covers the period 1991q1-2009q3.

[9] Notice that the t_I test corresponds to the t-value of the error correction term in Equation (4), which is negative and significant according to the tabulated critical values in Pesaran et al. (2001).

[10] See Brouwer and Ericsson (1998), Banerjee and Russell (2001) and Bowdler and Jansen (2007).

[11] In order to avoid too many tables, we only provide the symmetric model together with the most suited asymmetric model (according to the Wald test for short and long run asymmetry). That is, the asymmetric model in the tables correspond either to the model with long-run symmetry and short run asymmetry, long run asymmetry and short symmetry or long-run and short asymmetry. We also present only the estimated coefficients of our variables of interest, namely the exchange rate. Results from the complete equations are available upon request from the authors.

Table 1: Bounds symmetric and asymmetric cointegration tests and Wald tests for symmetry

		LR & SR symmetry	LR symmetry & SR asymmetry	LR asymmetry & SR symmetry	LR & SR asymetry
Canada	t_I	-4.28^a	-3.10^w	-5.23^a	-5.34^a
	F_I	6.31^a	6.91^a	8.49^a	8.66^a
	W_{LR}			17.95^r	16.59^r
	W_{SR}		r		r
France	t_I	-4.61^a	-4.76^a	-3.83^w	-3.65^w
	F_I	8.46^a	10.15^a	7.47^a	7.22^a
	W_{LR}			20.38^r	20.38^r
	W_{SR}		0.50^{nr}		2.36^{nr}
Germany	t_I	-5.83^a	-4.99^a	-4.28^a	-6.00^a
	F_I	11.58^a	10.89^a	11.03^a	12.36^a
	W_{LR}			10.36^r	16.44^r
	W_{SR}		r		11.04^r
Italy	t_I	-4.59^a	-4.58^a	-4.56^a	-5.14^a
	F_I	8.78^a	8.43^a	8.18^a	15.71^a
	W_{LR}			0.12^{nr}	0.71^{nr}
	W_{SR}		r		r
Japan	t_I	-5.72^a	-5.72^a	-6.67^a	-5.64^a
	F_I	9.16^a	9.16^a	9.37^a	8.83^a
	W_{LR}			3.57^r	1.36^{nr}
	W_{SR}		r		r
USA	t	-4.73^a	-3.18^w	4.57^a	-3.34^w
	F_I	6.20^a	8.72^a	6.24^a	6.94^a
	W_{LR}			29.58^r	33.92^r
	W_{SR}		r		r
UK	t_I	-6.58^a	-5.97^a	-8.08^a	-6.25^a
	F_I	11.01^a	7.90^a	11.94^a	8.95^a
	W_{LR}			5.34^r	7.87^r
	W_{SR}		r		r

Notes: (1) t is the t-ratio for testing $H_0 : \pi_p = 0$ in equation (??), (4); (2) F is the F-statistic for testing $H_0 : \pi_p = 0$ and $H_0 : \pi_s = 0$ and $H_0 : \pi_Y = 0$ in equation (4); (3) W_{LR} refers to the Wald test of long-run symmetry; (4) W_{SR} denotes the Wald test of the additive short-run symmetry condition; (5) ^{b} indicates that the statistic lies below the 0.005 lower bound (6) ^{w} the statistic falls within the 0.005 bounds; (5) ^{a} it lies above the 5% upper bound; (7) ^{r} denotes rejection of symmetry; (8)^{nr} denotes not rejection of symmetry; (9) Critical values from Pesaran et al. (2001); (9) we follow the general-to-specific approach to select the final ARDL specifications, starting with a maximum lag length of 4 and then dropping all insignificant stationary regressors.

and the exchange rate, indicating that depreciations (i.e. when s_t rises) is associated with an increase in the consumer prices, as predicted by the law of one price and the PPP theory.
However, it is interesting to notice that in several cases, allowing for asymmetry increases both the magnitude and the significance of the pass-through. Indeed, in the restricted symmetric model, the estimated long-run coefficients of the exchange rate (L_s) for France, Germany and the USA are not statistically significant. This result is consistent with several studies which do not find a cointegrating relationship between the exchange rate and the prices in these countries[12]. Yet, once asymmetry -in the short, in the long or in both terms- is allowed for, the exchange rate enters the long run relationship as a significant component in the three countries (at least one of the long-run partial sums coefficients as we detail below). This first important result underscores the importance of accurately specifying the long-run relationship. It suggests that the lack of evidence in the previous studies may be due to the failure to correctly model the asymmetric relationship.
In fact, results in tables (2) to (3) suggest strong evidence of asymmetry although a great heterogeneity among countries. Remember that micro-oriented works using disaggregated data find evidence of asymmetric responses but no clear pattern as the direction changes accross different industries. Evidence of asymmetry from aggregate data is thus a novel interesting result.
For instance, consider the case of the USA and Germany, the results indicate that prices increase rapidly and strongly following a weakening of the domestic curreny but the response to an appreciation is distinctly milder. This asymmetric effect is confirmed by the Wald tests, W_{LR} test statistics in table (1), which clearly rejects long-run linearity. These results suggest that the US and Germany import markets are dominated by an imperfect competition structure which allows the foreign exporters to resort to price to market.
Bussière (2007) finds a similar result for Germany in the short-run dynamics: large appreciations have a less than proportional effect. Germany which is the largest exporter in the European area is also a majour importer of intermediate imputs. The German industrial imports are high-quality sophisticated and specialized inputs used in their own export sector. This may explain an oligopolistic position of foreign exporters who can afford a lower pass-through after a devaluation than after an appreciation. In addition some quantity constraints are highly probable in chemical and petroleum related industries, providing inputs to Germany. Finally this result suggests a low substituability between local and imported inputs in the German production structure.
On the other hand, our result on the US are consistent with Swammy and Thurman (1994) who also found a higher pass-through with depreciations in the US using industrial data. This

[12] Two reference studies in the markup literature among other fail to find a cointegrating relationship bewteen import prices and the prices. They are by Franz and Gordon (1993) who do not find a significant relationship between the exchange rate and prices in the USA. And the study by Banerjee and Russel (2001) on G7 economies, where import prices enter the markup with an insignificant coefficient for Japan, Germany, France, and Canada.

result suggests that there is a low substituability between local and imported goods in the US. This last feature may result from the structural transition of the US production from industrial to service goods. In the case of the US, evidence of asymmetry from aggregate data is a novel interesting result.

Last, note that the US and Germany have a similar degree of pass-through included between 9 and 16%. It means that a depreciation by 10% is associated with an increase in the consumer price level by 0.9 and 1.6%. On the other hand, according to our results an appreciation is not transmitted to the prices.

On the contrary, Canada, France and the UK are characterized by completely the opposite effect: in the long-run, whereas depreciations do not have significant effects, appreciations of the local currency are met by important reductions in prices. Once again, the long-run Wald tests, W_{LR}, are able to firmly reject the null hypothesis of long-run symmetry in all cases. While Canada and the UK have a similar ERPT degree o around 10%, France features a surprisingly high pass-through of 80% after an appreciation. The asymmetry result in the UK contradicts the finding of Herzberg, Kapetanios and Price (2003) who found no evidence for non-linearity during a similar period as ours[13]. In total, the asymmetry results suggest a more competitive context in these countries and/ or a lower dependence to imports due to a higher substituability between domestic and imported goods than in the US and Germany.

In the case of Italy and Japan the long-run coefficients for the positive and negative partial sums are significantly similar when both long-run and short-run asymmetry is imposed, reflecting a linear and symmetric effect of nominal appreciation and depreciation upon inflation. Yet, in the case of Japan, given that the short-run partial sums are not significant, our preferred model imposes symmetry in the short-run (model III). In this case, symmetry is rejected by the Wald test, reflecting the fact that prices are reduced slightly if the Yen appreciates but are less passed through to higher prices after a depreciation. This result is consistent with the difficulty to enter the Japanese market for foreign exporters due to strong entry barriers. Hysteresis effects emphasized by Froot and Klemperer (1989)implying that a temporary loss in market share is likely to become permanent amplify asymmetries. Our results suggest thus that foreign exporter firms attempting to increase or maintain their market share in Japan increase pass-through when the importer's currency is appreciating and decrease pass-through when the importer's currency is depreciating.

Turning now to the analysis of short-run, we found that only a few lags are enough to capture the dynamic adjustment both when symmetry is imposed and when the models allows for short-run asymmetry (either with long-run symmetry or asymmetry). However, in light of the overwhelming rejection of long-run symmetry in all cases, the preferred models are therefore either with long-run asymmetry & short-run symmetry or long-run & short-run asymmetry. In 3 out of the 7 countries, short-run asymmetric dynamics seem to be important, either because just lags of positive partial sums are positive and significant, as it is the case in

[13]However they acknowledge that their tests are not very powerful in samples of the size they consider.

Canada and Italy; or, like in Germany, because when both partial sums are significant, the Wald test for short-run dynamics rejects the null of symmetry (see W_{SR} in table (2). In the rest of the countries, since the partial sums in first differences are not significant, there is little evidence of short-run asymmetry. For these countries, the relevant model imposes therefore only asymmetry in the long-run.

In the short-term relationship, only positive changes in the exchange rate (depreciation) are significant determinants of inflation in Canada. An increase in the depreciation rate by 10% increases inflation by 0,5% while no adjustment takes place in the case of an appreciation. Inflation is determined by the variation of the depreciation rate in these three countries only. Short-run asymmetric dynamics can be well interpreted within a graphical analysis of the dynamic multipliers.

In conclusion our results suggest strong evidence of asymmetry and a great heterogeneity among countries. Such asymmetries imply asymmetric adjustment of a balance of payments after a depreciation and after an appreciation. The adjustment differ across the countries. We come back to this implication in our conclusion.

Table 2: **Dynamic symmetric and asymmetric estimation of price adjustments: Canada, France, Germany and Italy**

	Symmetric			Asymmetryic		
	Variable	Coeff.	t-stat	Variable	Coeff.	t-stat
Canada	p_{t-1}	-0.046	-4.28	p_{t-1}	-0.110	-5.34
	s_{t-1}	0.012	2.93	s^+_{t-1}	-0.010	-1.24
				s^-_{t-1}	0.016	3.63
	Δs_t	-0.034	-2.19	Δs^+_{t-1}	0.050	2.00
	Δs_{t-1}	0.042	2.69			
	L_s	0.266	3.43	L^+_s	-0.091	-1.24
				L^-_s	0.142	3.83
France	p_{t-1}	-0.018	-4.61	p_{t-1}	-0.037	-3.83
	s_{t-1}	-0.003	-0.61	s^+_{t-1}	-0.016	-2.00
				s^-_{t-1}	0.033	3.24
	Δs_t	0.039	2.46	Δs_t	0.032	1.99
	Δs_{t-3}	0.036	2.24			
	L_s	-0.180	-0.58	L^+_s	-0.439	-1.64
				L^-_s	0.898	6.28
Germany	p_{t-1}	-0.065	-5.83	p_{t-1}	-0.169	-6.00
	s_{t-1}	0.008	1.07	s^+_{t-1}	0.027	3.17
				s^-_{t-1}	-0.027	-1.28
	Δs_{t-i}	n.s	n.s	Δs^+_{t-3}	0.160	3.13
				Δs^-_{t-3}	-0.099	-2.17
	L_s	1.09	0.120	L^+_s	0.162	3.23
				L^-_s	-0.162	-1.48
Italy	p_{t-1}	-0.053	-4.59	p_{t-1}	-0.053	-4.58
	s_{t-1}	0.036	5.22	s_{t-1}	0.035	4.95
	Δs_t	0.105	5.40	Δs^+_t	0.123	5.28
	Δs_{t-1}	-0.049	-2.46	Δs^+_{t-1}	-0.062	-2.54
	L_s	0.689	5.01	L_s	0.658	4.98

Notes: (1) L_s, L^+_s and L^-_s are the general long-run coefficient associated with the general, positive or negative changes of exchange rates or import prices, respectively

6 Conclusions

In this paper, we have followed a novel and simple method of combining asymmetric cointegration with a dynamically flexible ARDL model in order to explore exchange rate pass-through in the G7 countries over the period 1970q1-2009q3.

By taking into account the asymmetric effects of nominal depreciation and appreciation on

Table 3: **Dynamic symmetric and asymmetric estimation of price adjustments: Japan, United Kingdom and United states**

	Symmetric			Asymmetric		
	Variable	Coeff.	t-stat	Variable	Coeff.	t-stat
Japan	p_{t-1}	-0.216	-5.72	p_{t-1}	-0.288	-6.67
	s_{t-1}	0.016	3.81	s^{+}_{t-1}	0.013	2.71
				s^{-}_{t-1}	0.020	4.33
	Δs_{t-i}	n.s	n.s	Δs_{t-1}	-0.022	-2.13
	L_s	0.075	4.18	L^{+}_{s}	0.047	2.72
				L^{-}_{s}	0.069	4.08
UK	p_{t-1}	-0.211	-6.58	p_{t-1}	-0.250	-8.08
	s_{t-1}	0.025	3.81	s^{+}_{t-1}	0.002	0.25
				s^{-}_{t-1}	0.025	3.54
	Δs_{t-i}	n.s	n.s	Δs_{t}	-0.041	-2.34
				Δs_{t-4}	-0.042	-2.24
	L_s	0.118	3.67	L^{+}_{s}	0.010	0.25
				L^{-}_{s}	0.100	3.48
USA	p_{t-1}	-0.032	-4.73	p_{t-1}	-0.067	-4.57
	s_{t-1}	0.005	1.84	s^{+}_{t-1}	0.007	2.11
				s^{-}_{t-1}	-0.005	-0.82
	Δs_{t}	0.022	2.01	Δs_{t-4}	0.001	2.05
	L_s	0.1711	1.90	L^{+}_{s}	0.0925	2.06
				L^{-}_{s}	-0.090	-1.40

Notes: (1) IDEM table (2)

prices, we go further than previous research which adheres to a linear paradigm, reflecting the assumption that positive and negative variations of the exchange rate have symmetrical effects on prices. However, as shown by our results, neither the price level nor the inflation rate does not behave in this naive way.

Indeed, based on the asymmetric ARDL approach, the following conclusions can be made. First, there is clearly a role for the exchange rate in the determination of the price in the long-run, even if exchange rate changes are less than completely associated with changes in the consumer price. Our modelling strategy emphasizes a cointegrating relationship between the exchange rate and prices in all G7 countries, contrary to several past studies. This first important result underscores the importance of accurately specifying the long-run relationship. The lack of evidence in the previous studies may be due to the failure to correctly model the asymmetric dynamics.

At the same time, the responsiveness of inflation to exchange rate variations is neither linear nor homogeneous accross countries. In particular, there seems to be downward rigidities in the USA and Germany which result in a lower pass-through measured after an appreciation than after a devaluation. The low substitutability between imported and local goods may be a key determinant for this asymmetry in these two countries. Germany is a large exporter, highly dependent from imported inputs. The industrial structure is specialized in the final export goods and substitute to imported goods are scarece. On the contrary, in Canada, France, the United Kingdom and Japan, domestic prices do not respond much to exchange rate depreciations in the long-run. In these countries, we find that the pass-through is higher in the case of an appreciation than a depreciation. In Japan this result is consistent with the difficulty to enter the domestic market for foreign exporters due to strong entry barriers. Foreign exporter firms attempting to increase or maintain their market share in Japan increase pass-through when the importer's currency is appreciating and decrease passthrough when the importer's currency is depreciating. Nonetheless, overall, our results confirm that, in the short-run, prices react indistinctly to depreciations than to appreciations.

In addition, our results suggest a heterogenous adjustment of the balance of payment accross the G7 countries. The fact that the pass-through is only significant after a depreciation in a country implies that a depreciation result in inflation and a more limited variation of the real exchange rate than after an appreciation. This dynamics thus reduced the impetus for a balance of payment adjustment after a devaluation. Given that this dynamic is evidenced in the United States and Germany has different implications. While a fast adjustment of the balance of payment is required in the United States in the aftermath of the financial crisis, the asymmetry results suggest that it will take longer than in other developed countries. Our results on Germany highlight the counter-effective effects of a devaluation of the euro currency on the German competitiveness: a devaluation of the Euro increases inflation in Germany more than an appreciation reduces it, which in fact deteriorates its competitiveness. This result is a bad news for Europe with Germany being the engine of Western European exports.

Finally, falling to account for this asymmetry pass-through can have important consequences in the sense that not taking into account the different effects of appreciations and depreciations, the real pass-through is not revealed. In this sense, the estimated importance of the asymmetric effect exerted by the nominal exchange rate may pose important dilemmas for policy-makers wishing to achieve both price stability and export competitiveness. Indeed, an important policy implication of these findings is that the different dependence of the exchange rate pass-through on positive and negative changes should be taken into account in deciding monetary policy rules.

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