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Exchange Rate Movements and Export Prices an Empirical Analysis

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Thanks to seminar participants at CREST for stimulating questions. I am especially grateful to E. Duguet, Y. Yoshida and G. Laroque for their advices.

Exchange rate movements and export prices an empirical analysis

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Abstract

This paper studies the effect of changes in the exchange rates on international trade prices by estimating exchange rate pass-through elasticities for the 500 largest exporting industries of six OECD countries. In around half of the industries, export prices appear sensitive to exchange rate movements, casting doubt on the assumption of a full pass-through of exchange rate movements into import prices at the root of the Purchasing Power Parity relation. The high heterogeneity of the results both between sectors and between national exporters suggests a pre-eminent role for structural factors, rather than macroeconomic ones in the explanation of the phenomenon.

Résumé

Ce papier étudie l'effet des mouvements de change sur les prix du commerce international en estimant l'élasticité des prix aux variations de change dans les 500 plus grosses industries exportatrices de 6 pays de l'OCDE. Dans environ 50% des secteurs, les prix à l'exportation apparaissent sensibles aux mouvements de change, en contradiction avec les hypothèses à la base de la relation de parité des pouvoirs d'achat. La forte hétérogénéité des coefficients, à la fois entre secteurs et entre exportateurs de différentes nationalités, suggère que l'explication de ce phénomène est plus structurelle que liée à des facteurs de type macroéconomique.

Keywords: exchange rate pass-through, macroeconomic environment, price strategies, panel data
J.E.L. Classification: F14, F41

1 Introduction

The sensitivity of prices to exchange rate movements is a recent but fruitful subject in Open Macroeconomics debates. Indeed, since Betts and Devereux [1996] model of pricing-to-market, the hypothesis of a full pass-through of currency changes into import prices, at the root of the Law of One Price, has been questioned many times. Empirically, recent papers have put in evidence an imperfect elasticity of aggregate import prices to exchange rate movements in numerous countries (see Campa and Goldberg [2002] for a cross-country comparison). From a theoretical point-of-view, the supposed degree of exchange rate pass-through (ERPT hereafter) has been shown to be an important determinant of both the international monetary transmission and the determination of the equilibrium exchange rate. However, if the macroeconomic consequences of the incomplete pass-through are well-documented, the explanation of the underlying nominal deviations is not yet the subject of a consensus. More precisely, two kinds of explanations coexist, either the exporters face nominal rigidities (as in Betts and Devereux [1996] paper [1]) or there are price discriminating strategies, rationalized in some environments (as in the Pricing-to-Market literature, following Krugman [1986]).

The objective of this paper is to take part in this debate from an empirical point-of-view. Indeed, the applied literature has often neglected the supply-side aspect of the price determination to focus on the macroeconomic consequences of the Incomplete pass-through (see Devereux and Yetman [2002], Anderton [2003], Herzberg, Kapetanios and Price [2003] for recent contributions). Microeconomic evidence is only provided by Industrial economics studies (see Mann [1986], Gagnon and Knetter [1991], Knetter [1989, 1992, 1993], Gil-Pareja [2003], etc.) that however focus on some specific sectors, leading to results that cannot be generalized easily. Yet, a microeconomic study could provide insights on the origin of the disparity of pass-through elasticities observed in cross-country comparisons. For instance, Campa and Goldberg [2002] estimate the long-run pass-through into import prices to vary between 40 and 130 % depending on the country. To explain those disparities, two kinds of arguments are conceivable : those based on specific macroeconomic features of the countries (as in Taylor [2000] model for instance) and those invoking industry-specific pass-through strategies (as in the Pricing-to-Market literature) leading to aggregate disparities when countries do not import identical baskets of goods. The present paper provides elements in

1. Betts and Devereux's main hypothesis is that some exporters set their price directly in the buyer's currency, an invoicing strategy that leads, when prices are rigid, to a zero pass-through of exchange rate movements into import prices. For attempts to rationalize such strategies, see Bacchetta and van Wincoop [2004] and Devereux and Engel [2001].

this debate, by studying the sensitivity of export prices to currency changes[2] on a large set of industries located in six exporting countries. When compared to aggregate estimates, the main benefit of this industry-based method is to weaken the twofold risk of endogeneity and composition bias[3]. Since covering a large part of the countries' exports[4], the estimates can be compared in several dimensions to ask i) if the incomplete pass-through observed at the aggregate level is also apparent in sectorial data and, if the case arises, ii) if it is common to all sectors (in which case macroeconomic features are more likely to explain it) or, on the contrary, if aggregate results are mainly driven by a few industries (in which case one has to focus on structural determinants to rationalize local incomplete pass-through strategies).

Results provide support for the incomplete pass-through hypothesis. Indeed, even if a large part of the estimated pass-through coefficients crowds around zero (a value that is compatible with complete ERPT), the share of coefficients that are significantly different from zero is too large to be due to a simple statistic phenomenon, thus reflecting a real incomplete pass-through phenomenon. Moreover, the asymmetry of distributions towards negative values is consistent with the Pricing-to-Market model : in those industries, the exporters' response to exchange rate movements leads to a mark-up adjustment that stabilizes local prices. Because this kind of strategy is likely to be influenced by features of the destination market, some importer-specific pass-through coefficients are then estimated, that however fail to show a systematic differentiation of pass-through strategies among importers. In some cases, importing countries seem effectively subject to a more pronounced price discrimination but, as this tendency is limited to a small number of industries, one cannot give it a macroeconomic interpretation. When comparing exporter-specific distributions, some slight differences also appear. In particular, the strong asymmetry of the Italian distribution towards negative values suggests a higher propensity to price-to-market from Italian exporters. Such national features must however be interpreted cautiously. Indeed, the high heterogeneity of estimates, both between industries and between expor-

2. Because I am working on export rather than import prices, the benchmark for the pass-through elasticities is different from that of the standard related literature. Indeed, whereas the pass-through is said to be complete when the ERPT elasticity of import prices (expressed in the consumer's currency) is unitary, here the Exchange Rate Pass-Through is complete when the elasticity of export prices is zero, meaning that export prices are insensitive to currency changes.

3. Indeed, Campa and Goldberg [2002] underscore the presence of composition bias in pass-through estimates based on price indexes. The endogeneity question is also brought up in several papers.

4. By contrast with the previously cited sector-based estimations that focus on a limited number of industries...

ters in a given sector, suggests a pre-eminent role for firm-specific factors to explain incomplete pass-through strategies. The higher mean propensity of Italian exporters to price-to-market could thus be explained in several ways, such as by its pattern of specialization, by specificities in its market structures or by its macroeconomic environment.

The model, the empirical method and the data are presented in section 2. In the third section, I present the results obtained under the assumption of an homogeneous pass-through in the whole sample of importing countries. The sensitivity of the results to the homogeneity assumption is then studied. Last, section 4 concludes.

2 Estimation of Pass-Through coefficients

2.1 The Model

As previously announced, this paper presents systematic estimations of ERPT elasticities for a large set of industries using a simple model. Each industry in a given country is assimilated to a representative firm that sells its goods in N separated national markets ($j = 1...N$) and chooses its prices by solving N maximizing problems of the following form :

$$\max \quad \Pi_t^j = Max \quad [P_t^j D_t^j - C(D_t^j)]$$

where :

- P_t^j is the optimal price for sales in the country j (supposed to be denominated in the exporter's currency),
- D_t^j is the demand addressed by country j to the considered producer.
- $C(D_t^j)$ is the total cost of producing D_t^j,

If the technology is a constant returns one, the marginal cost (Cm_t) is independent from the sold quantities (and of the destination country). First-order conditions then lead to a set of N equations :

$$P_t^j = \mu_t^j Cm_t, \quad j = 1,...,N \tag{1}$$

where μ_t^j, the optimal mark-up, depends on the perceived elasticity of demand η_t^j :

$$\begin{aligned} \mu_t^j &= \frac{\eta_t^j}{\eta_t^j - 1} \\ \eta_t^j &= -\frac{\partial D_t^j / D_t^j}{\partial (P_t^j S_t^j)/(P_t^j S_t^j)} \quad = \quad \eta^j(P_t^j S_t^j, Z_t^j) \end{aligned}$$

Optimal prices are thus a function of the production costs (Cm_t) and the elasticity of demand with respect to the local price ($P_t^j S_t^j$ with S_t^j the nominal exchange rate expressed in terms of j's money per unit of the exporting country's currency). To keep things as general as possible, this elasticity is itself written as a function of the consumer price and of a set of demand attributes (Z_t^j). Using a Taylor approximation on the log-linearized form of (1) leads to a theoretical relation that links (locally) the export price to the marginal cost, the features of demand and the exchange rate[5]:

$$p_t^j = \alpha^j + (1+\beta^j)cm_t^j + \frac{\varepsilon_Z^\eta}{\varepsilon_{PS}^\eta}\beta^j z_t^j + \beta^j s_t^j \tag{2}$$

with lower cases used for the logarithms of the corresponding variables in levels and:

$$\alpha^j = (1+\beta^j)\ln\mu_0^j, \quad \beta^j = -\frac{\varepsilon_{PS}^\eta}{\eta^j - 1 + \varepsilon_{PS}^\eta}, \quad \varepsilon_{PS}^\eta = \frac{\partial \ln \eta_t^j}{\partial(p_t^j + s_t^j)}, \quad \varepsilon_Z^\eta = \frac{\partial \ln \eta_t^j}{\partial z_t^j}$$

In (2), β^j measures the sensitivity of export prices to exchange rate movements, the "pass-through coefficient" specific to each destination country. As demonstrated in the Pricing-to-Market literature (see Krugman [1986]), this elasticity depends on the convexity of the perceived demand schedule (reflected in ε_{PS}^η): it is null when the elasticity of demand with respect to the consumer's price is constant ($\varepsilon_{PS}^\eta = 0$), in which case the optimal pass-through is complete, negative for a less convex demand schedule, when the perceived elasticity of demand is an increasing function of the local currency price ($\varepsilon_{PS}^\eta > 0$)[6] and positive in the opposite situation. Following Krugman, numerous papers studied the optimal pass-through strategies in some specific analytical frameworks and competitive environments and the influence of a number of determinants. One can classify them into three main categories: i) those that are firm-specific, as the importance of intra-firm trade in the exporter's total sales (Rangan et Lawrence [1993]), ii) industry-specific determinants, as the number of firms established in the destination market (see e.g. Dornbush [1987]), the size of entry costs in the importing country (Dixit [1989]) or the degree of price rigidity (Kasa [1992]), iii) and macroeconomic influences as the bilateral exchange rate volatility or the monetary environment in the destination market (Devereux and Engel [2001]). With a possible combination of several of those aspects, the exact degree of pass-through is difficult to anticipate; the only certainty is that, in a segmented world, the hypothesis of a full pass-through, at the root of the standard Law of One Price, can be invalid in numerous situations.

5. See appendix A1 for details

6. Indeed, with a convex demand schedule, a depreciation of j's currency (an increase in S_t^j) leads to a decrease in the demand that must be compensated by a cut in the margin. The opposite is true when the currency appreciates.

2.2 Empirical specification

The estimation equation used to identify the pass-through coefficients β^j is directly obtained from (2) :

$$p_t^j = \alpha_0^j + \alpha_1^j cm_t + \alpha_2^j z_t^j + \beta^j s_t^j + \varepsilon_t^j \tag{3}$$

In this equation, the parameter of interest is β^j, the elasticity of prices to exchange rate movements. On the other hand, both α_1^j and α_2^j are difficult to estimate with accuracy because of measurement problems, linked to the unobservable nature of marginal costs and the undetermination of variables included in Z_t^j without an exact definition of the functional form of demand (D_t^j). Moreover, beyond those measurement problems, a more serious risk arises due to the potential correlation between those explanatory variables and the exchange rate. For instance, when part of the inputs are imported from abroad, the marginal cost is influenced by exchange rate movements. In this case, an exact estimation of (3) requires precise assumptions on the cost structure (see Athukorala and Menon [1994]).

To keep things more general, one has to turn to a statistical version of (3). The approach in this paper is in the line of Knetter [1989], that uses the multidimensional variability of time-series data on export prices towards a cross-section of destinations. The idea is the following : if the exporters costs (cm_t in (3)) are independent of the country of destination a complete set of time dummies incorporated in (3) instead of a (probably mismeasured) marginal cost variable will identify them rather well in a panel framework. In the same way, one can hope to capture part of the country-specific features leading to price adjustments towards a specific market (z_t^j) with a complete set of individual effects. The present paper then estimates exchange rate pass-through elasticities on distinct samples of industry-specific export prices towards several countries in a given period, using a model with time and individual fixed effects to control for a large array of cost and demand shocks :

$$p_t^j = \alpha_t + \gamma^j + \beta^j s_t^j + \varepsilon_t^j \tag{4}$$

In (4), the residual term ε_t^j will then catch price adjustments that are neither linked to exchange rate movements, neither due to constant country-specific features, neither common to all the importers. For instance, part of punctual country-specific shocks could pass in this error. In the estimates, those residuals are supposed to be *i.i.d.* for technical convenience but, in some sectors, serial or spatial correlation is able to bias results, in which case a full analysis would require a specific correction. Ideally, one should also estimate (4) with an instrumental variables method, since part of the shocks to export prices might affect the exchange rate as well. It is however not clear how best to instrument this variable. Moreover, because of the descriptive approach of this work, the estimation method must be simple enough to be used systematically on a large number of industry-specific samples. In the following, the hypothesis of orthogonality between exchange rates and the residual is maintained.

Asymptotically efficient estimates of the coefficients in (4) are obtained using a Feasible Generalized Least Squares method that supposes randomly distributed individual effects (γ^j). The estimated equation thus takes the following form :

$$p_t^j - (1-\hat{\theta})\bar{p^j} = \alpha_t + \beta^j(s_t^j - (1-\hat{\theta})\bar{s^j}) + \epsilon_t^j$$

where :

$\bar{x^j} = \frac{1}{T_j}\sum_{t=1}^{T_j} X_t^j$ is the averaged value of the variable X, computed for each individual on its period of presence in the sample (that is specific to each individual in the unbalanced panel),

and $\hat{\theta} = \frac{\hat{\sigma_W}}{\hat{\sigma_B}}$ is the ratio of estimated variances obtained by estimating successively the within and the between form of (4) [7].

A drawback of the FGLS method is that it is not consistent if the individual effects are correlated with explanatory variables. Since the individual effects in (4) are used as a proxy for country-specific features affecting the price behaviours, one can not definitely exclude the possibility of a correlation with the exchange rate. To check the consistency of the estimation, a Hausman test (where the null hypothesis is the absence of correlation between the fixed effects and the explanatory variables) is computed at the end of the estimation. When the null hypothesis of no correlation is rejected at the 5% level, the pass-through estimate that is kept is the one obtained by applying the Ordinary Least Squares to the within form of (4).

Even with those transformations allowing to avoid the need to estimate the individual effects, the number of coefficients to be estimated is often too high as long as ERPT coefficients are authorized to be different for each individual (for each destination market). Following Knetter [1993] [8], the estimation is thus done assuming that the pass-through coefficient is the same whatever the importing country ($\beta^j = \beta$, $\forall j$). This assumption can at least be interpreted as providing mean pass-through coefficients for each source-industry pairs, that are easier to interpret than a large number of industry-

7. See Appendix A.2 for details

8. Knetter [1993] uses the same type of model to estimate importer-specific pass-through coefficients and shows that the null hypothesis of identical values of coefficients across destinations is rejected in only 8 of its 52 source-industry pairs.

and destination-specific ERPT. In a second step however, the validity of this assumption is checked in estimations adding some country-specific exchange rate series.

2.3 Data

Equation (4) is estimated separately on samples covering the exports of six countries (Germany, the United States, France, Italy, Japan and the United Kingdom) towards OECD importers [9] in around 2000 industries at the 5-digit SITC commodity level. Data cover a period from 1988 to 1998 (1989-1998 for the USA) at an annual frequency that permits to limit the serial correlation (well-known to be serious in price equation because of nominal rigidities). Indeed, assuming that most of the price adjustments occurs in the first months, it is not necessary to add a lag in the estimated equation. From a technical point-of-view, the estimation is simpler since the control of serial correlation in panel data would require a GMM estimation. The main advantage of annual data is however more fundamental. Indeed, since the objective of this paper is to study the pass-through behaviour from the producer's point-of-view, it is convenient to estimate a coefficient that is not influenced by short-run factors, such as nominal rigidities, but actually reflects the firm's strategic pass-through decision. In their estimation using quarterly data, Campa and Goldberg [2002] assimilate the long-run pass-through to a combination of coefficients on 4 lags of the exchange rate. Based on this result, one can think of the ERPT coefficient obtained from annual data as a long-run coefficient, that reflects strategic behaviours rather than nominal constraints. The problem of serial correlation could however still exist in industries in which menu costs are so large that price adjustments spread over more than one year...

The previous panel specification only requires two data series, namely export prices and exchange rates. Because of a lack of disaggregated bilateral price data, one has to proxy export prices by F.O.B. [10] export unit-values (i.e. quotients of values by quantities) obtained from the OECD's *International Trade by Commodities Statistics* database. As well-documented by the

9. Importers (j) are restricted to OECD countries, in order to have long enough series and sufficient traded volumes. Moreover, the assumption of an homogeneous pass-through whatever the destination country is more likely to hold between countries whose economic structures are nearby.

10. F.O.B.(*Franco On Board*) data do not include transportation, tariff wedges and local costs. The use of such data is then consistent with our initial specification that does not take into account transport costs. This treatment still is questionable if transport costs influence pricing strategies or are correlated with exchange rate movements...

trade literature (see e.g., Kravis and Lipsey [1974]), measurement errors in this proxied dependent variable are then a potential source of disturbance in the regression, leading to an upward bias if correlated with exchange rate movements. Those measurement errors are however more salient with quarterly data (because of the high volatility of volumes due to consumption transfers) and at a higher aggregation level [11]. Limiting the aggregation bias thus motivates the choice of a 5-digit SITC nomenclature. However, since the measurement problem still exists and is more pronounced when computing unit-values on a small volume of transactions, the discussion is limited to the 500 largest exporting sectors of each country. This *ad hoc* selection permits to increase substantially the accuracy of the estimated coefficients (as measured by the mean estimated standard error) with a limited loss of generality since the concerned industries still cover more than 70% of the total value of exports, whatever the exporting country (and even more than 90% of Japanese exports). Moreover, this selection does not forbid a structural analysis since industries from each 1-digit classes of the SITC nomenclature are represented in the limited sub-sample, except for Japan and the United Kingdom [12]. The 7th class (Machinery and transport equipment) is however obviously dominant in the exporting activity of the countries under study, with between 27 and 48 % of the sampled industries in it.

As far as the exchange rate is concerned and following Knetter [1989], two alternative series have been used : the nominal exchange rate between the exporting country and each of its partners [13] and the same exchange rate adjusted by the importer's price level (measured by the CPI). By this correction, one estimates the sensitivity of prices to real exchange rate movements (price adjustments that are not attributable to the general inflation in each market). Adjusting for the exporter's CPI would not make any difference because of the presence of time dummies and the logarithmic form of the equation (see Takagi and Yoshida [2001]). Figure 1 and statistics in Table 1 show the evolution of the nominal exchange rates (expressed as the number of units equivalent to 1 dollar). In most cases, the exchange rate fluctuates slightly around a more or less stable value. The currency of the less develo-

11. Indeed, the more data are aggregated, the more heterogeneous are industries, leading to quality differences between the goods produced by the "representative firm". In that case, the use of unit values can be dangerous because of the implied price differences that averages are unable to reflect. Unit values are then, *a priori*, a more accurate proxy of disaggregated prices.

12. No industries from the 1st (Beverages and tobacco), 3rd (Mineral fuels, lubricants and related materials) and 4th (Animal and vegetable oils, fats and waxes) classes belongs to the 500 largest exporting industries in Japan, neither do industries from the 4th class in the United Kingdom.

13. With all bilateral exchange rates normalized to one dollar in 1995.

ped countries (Turkey, Mexico, Korea, Poland, Hungary) with regards to the dollar however depreciates continually throughout the period of estimation. Even when restricting the sample to OECD importing countries then, one can still study pass-through behaviours in various monetary contexts.

3 Results

3.1 Interpretation of crude results

Estimating systematically a high number of coefficients using such a basic model is obviously a dangerous game. However, restricting the sample to the largest industries permits to increase the mean precision and to limit heterogeneity : on average, the estimated standard error of the pass-through coefficient in each exporting country is less than 0.05 when using the nominal exchange rate (see Table 2[14]). The accuracy is however less with real exchange rates (Table 3), casting some doubts on those results and leading me to favor an analysis based on nominal ERPT coefficients. The approach to interpret the estimated sector-specific elasticities in terms of pass-through decisions from is to assimilate them to a sample of 500 pass-through coefficients drawn from the unknown "true" distribution, that gives us indications on this distribution properties. In view of those estimated distributions (illustrated in Figure 2), several questions come to mind : do they point to incomplete pass-through behaviors ? If so, are those effects consistent with the theoretical arguments of the related literature ? Are the properties of those distributions identical across exporters ?

Consider first the question of the actual existence of an incomplete pass-through phenomenon. Under the null hypothesis of a complete pass-through and using the 5% confidence level, the fraction of coefficients significantly different from zero in the drawn sample should be equal to 5 %[15]. Here, this

14. The statistics in tables 2-9 are simple averages of each distribution of sector-specific coefficients. Those that are interested in the macroeconomic implications of the estimates could wish to see weighted means rather than simple ones. Using a weighting scheme based on exported values, I have systematically calculated those statistics. The general picture is the same as with simple averages, probably because of the limitation to the largest industries. However, since I am mainly interested in the microeconomic aspect of the pass-through decision, I have chosen to keep simple means, that give the same weight to the behaviour of each type of exporters (large ones as well as smaller). Indeed, a systematic difference in the pass-through decisions of small producers would also provide insights on the determinants of pass-through strategies that would be undetectable with weighted averages.

15. Indeed, remind that a null coefficient β means that export prices are not sensitive to exchange rate movements, that the pass-through is complete.

fraction varies between 40 and 65% depending on the country and is thus too high to be interpreted as a statistical feature : the hypothesis of a systematic complete pass-through is rejected by the data, whatever the exporting country [16]. Those figures however still suggest that the "true" distribution of pass-through coefficients has a mode around zero, implying that the incomplete pass-through phenomenon is not generalized to the whole economy (as one could infer from aggregate estimates) but rather limited to around half of the industries.

On those significant coefficients, more than half are negative, reflecting an asymmetry of distributions that is consistent with the pricing-to-market explanation of the incomplete pass-through phenomenon. Negative coefficients indeed imply a partial absorption of exchange rate movements into the exporters margin in order to stabilize local prices, a strategy that is rational when the elasticity of demand increases with local prices and firms want to preserve their market shares. On the other hand however, the share of positive coefficients is too important to be explained by sampling errors, thus providing evidence of some pricing strategies that amplify the effect of exchange rate movements on import prices. The case for this kind of positive coefficients is generally ignored by theoretical papers since corresponding to highly convex demand functions that are difficult to rationalize (except for some specific goods). The evidence here however suggests that this analytical framework is not rich enough and should be completed to explain those strategies of local currency price amplification. One conceivable explanation might be related to the recent growth of intra-firm trade (see Rangan and Lawrence [1993] for evidence on pass-through behaviors in multinational firms).

When focusing on the distribution of the coefficients, as reflected by the interquartile ranges of the estimated coefficients (see the 4th line in tables 2 and 3), one sees that they are more concentrated around zero than one could have expected from the related literature. The fact that pass-through coefficients are higher (in absolute value) in previous industry-based studies is not surprising because of the already explained selection bias. However, since pass-through estimates obtained from aggregate prices show a very low pass-through of exchange rate movements into import prices [17], one would

16. The fractions of significant coefficients are relatively low compared to previous industry-based studies. This may be due to the fact that these studies use much more limited sample of industries, selected in an arbitrary way, either using previous evidences of pricing-to-market in some specific industries (as for instance in the automobile sector studied by Gagnon and Knetter [1995] and Gil-Pareja [2003] among others) or on the basis of structural features presented as being able to generate pricing-to-market behaviours (as in Knetter [1989] that chose to limit his work to homogeneous goods). Here, the only selection lies on a size criteria and the fraction of exports covered is large.

17. For instance, remember Campa and Goldberg's results : a long-run pass-through that

have expected to obtain more pronounced evidence of exchange rate smoothing. Here, most of the nominal ERPT coefficients lie between -0.3 and 0.2 [18], meaning that, in those industries, between 70 and 120% of the nominal changes are passed into export prices. The relative weakness of the estimated coefficients can be explained along two lines. First, one cannot rule out the possibility of a mispecification of the estimated equation, even if the presence of individual and time fixed effects controls for a large range of potentially omitted variables. In particular, the risk of serial correlation is not entirely removed for highly persistent nominal rigidities. In such a case, the omission of the lag of the price in equation (4) would create a bias on the estimated pass-through, that is positive as long as the current exchange rate and the lag of the price are positively correlated (see Appendix A3 for details). A large number of those biased estimates would thus increase artificially the mass of coefficients in the right part of the distributions in figure 2. However, part of the discrepancy could also be attributable to the non-standard perspective used in this paper, that leads me to consider export prices rather than import or consumer prices. Indeed, the sensitivity of estimates to the type of price series has been put in evidence many times : the pass-through is higher when estimated in import prices than in consumer (final) prices, suggesting that part of the incomplete pass-through phenomenon could be explained by the role of the distribution sector that intermediates between exporters and final consumers (see Tille [2000]). In the same way, differences in the magnitude of the ERPT estimated from import and export prices could reveal a sensitivity of this elasticity to trade costs, due to their correlation with the exchange rate. When assimilating trade costs to time-invariant unit transport costs, as standard in the trade literature, this possibility seems unlikely. However, a related literature has shown that the size of trade costs estimated indirectly from gravity equations is too high to be assimilated solely to transport costs (see the survey by Anderson and van Wincoop [2004]). Related to the present problem, one also can think of the cost of currency hedging, that is probably an important trade barrier and is of course not independent from exchange rate movements...

3.2 Multi-dimensional comparison of results

The comparison of exporter-specific distributions gives additional insights on the tendency of those countries exporters to adopt incomplete pass-through strategies. The German distribution is thus more concentrated around zero

varies between 40 and 130% according to the considered country.

18. Distributions are less concentrated with real exchange rates but part of the explanation is probably the lower accuracy of estimates...

than the five others means that, if any, the portion of exchange rate movements that is not passed into import prices is low. Figure 2 confirms this, putting in evidence a significant mass of coefficients between 0 and -0.2 that corresponds to strategies of weak local price stabilization. On the other hand, US and English distributions appear much more symmetric, meaning that explanations of incomplete pass-through in terms of nominal rigidities in the consumer's currency are not as convincing for those countries. The most striking feature of the comparison between exporting countries however lies in the strong asymmetry of the Italian distribution towards negative values : 15% of the considered estimates imply an absorption of exchange rate movements into mark-ups higher than 10%. On average, this country is the one whose tendency towards pricing-to-market strategies is the most pronounced.

Those exporter-specific results must however be taken cautiously. Indeed, the main feature of the distributions lies in the huge heterogeneity of results, implying large differences of exchange rate pass-throughs between 5-digit industries, that are smoothed in averaged statistics. Such a result is important since it introduces a doubt on the validity of results that rest on an homogeneity assumption, as do aggregate estimates. Even a crude distinction by types of products (as done for instance in Campa and Goldberg [2002]) will not necessarily remove the risk of composition bias since the heterogeneity is even observable within each class of products. Indeed, when studying the repartition of results between insignificant, positive and negative values for each 1-digit SITC class, results are far from conclusive, preventing us to single out a type of goods that would be particularly concerned by incomplete pass-through behaviours. For instance, whereas products from the 5th class (Chemicals and related products, n.e.s.) are over-represented in negative coefficients in US, German and Italian distributions, they are mainly positive in the English distribution : when estimations on US, German and Italian data suggest that the chemical industry is prone to pricing-to-market strategies, results are opposite with English exports. In this country, exporters appear rather liable to discriminate foreign markets in the 6th class (Manufactured goods classified chiefly by material) that is dominated by positive coefficients in the US and German distributions... Even at a finer level, I am unable to detect types of goods that exhibit a given pass-through behaviour in a recurrent way. For instance, when merging together the six sub-samples of coefficients that are significantly negative (that is to say the sub-samples of "pricing-to-market coefficients"), one obtains 627 industries on which only 267 appear in at least two exporter-specific distributions : more than half of the industries that are the subject of local currency pricing strategies by a given exporter are not in the five other countries. This disappointing result can be interpreted in two ways : either the SITC nomenclature does not put goods together in the convenient way as regards to the problematic[19], or the behavioral heterogeneity is too strong, linked to firm-specific rather than industry-specific determinants, in which case one would not be surprised by the little similarity of estimated coefficients at the sectorial level...

At this point of the analysis then, incomplete pass-through behaviours have been detected in a large number of sectors, in export data from the six considered countries, without however strong evidence on the kind of determinants that could explain those strategies. The only obvious element is that those determinants must be influential at the microeconomic level, explaining the absence of convincing recurrent features at the macroeconomic level. This finding is consistent with Bacchetta and van Wincoop [2004] theoretical model that puts in evidence several factors potentially influencing pass-through strategies, either macroeconomic (as the size of the exporting country), or sectorial (as the differentiation of goods) or firm-specific (as the market share) but concludes by favoring a structural explanation of the degree of pass-through, arguing that "the two main factors determining the invoicing choice are market share and differentiation of goods". If the differentiation of goods is the same whatever the location of the production, the market share is obviously exporter-specific, thus potentially explaining the high heterogeneity of results, even between exporters of the same industry located in different countries. To go further in the explanation of pass-through strategies, one should then work on a more structural model that would introduce explicitly those determinants. For this to be possible, a theoretical preliminary step is necessary : the building of a convincing framework that would reconcile the different explanations of pricing-to-market strategies in a general enough model to be tested in numerous industries. Of course, this is far beyond the scope of the present statistical analysis and will be left for further research...

Before concluding however, one has to check the validity of the hypothesis of an homogeneous pass-through behaviour towards all the considered partners. Since it is technically impossible to relax entirely this constraint, specific pass-through coefficients are estimated for groups of countries, that are chosen in view of results from previous empirical and theoretical papers.

19. Indeed, as argued by Davis et Weinstein (2003), industrial nomenclatures group goods with regards to their use rather than using technological criteria.

3.3 Robustness with respect to the sample of importing countries

To test the validity of the homogeneity assumption, this section presents results of additional estimations that add successively to equation (4) exchange rate series specific to some given sub-groups of importing countries. Results are then compared to those of the benchmark equation to verify if the addition of these variables does not change drastically previous results concerning the common pass-through coefficients β. Robustness checks are summarized in tables 4-7 that evaluate this sensitivity through the effect of the control on the results of the Hausman test (table 4), the significance, sign and range of the estimated common pass-through coefficients (tables 5, 6 and 7 respectively). In all the cases, one expects the results to be nearby those of the benchmark estimation, meaning that the hypothesis of an homogeneous pass-through does not bias things dramatically.

Beyond the sensitivity analysis, this exercise is also instructive by itself since it permits to evaluate the pass-through behaviour of an exporter towards a group of countries that share a common macroeconomic feature. Indeed, several papers suggest the possibility that exporters could have systematically different pass-through strategies towards some specific destination markets. In particular, the United States are often seen as being the subject of pronounced local pricing stabilization strategies, that could explain the low pass-through estimated from US import data. One explanation is that the US market is large enough for exporters to have a strong incentive to adjust their mark-up in order to maintain (or even to raise) their market share during periods of currency changes. To check the existence of such a tendency to stabilize US prices, the first added variable is a US-specific exchange rate series. The US-specific pass-through coefficient obtained in this way permits to ask i) if the pass-through strategy is significantively different towards the USA and towards the other OECD countries (as shown by the share of those coefficients that are significant in table 8), ii) if the direction of this differentiated effect is systematically the same throughout the sample of industries (as shown by the sign of the coefficients summarized in table 9).

The addition of a US-specific exchange rate series proves to have few effects on the common pass-through coefficients : results of the Hausman tests are the same as in the benchmark estimation in more than 80% of the industries, the sign of the mean pass-through is identical in almost all estimations and the interquartile ranges are remarkably similar (see line 1 in tables 4 and 5 and the comparison of lines 1 and 2 in table 7). Even if the share of significant coefficients slightly decreases, previous results seem robust to the addition of the US-specific explanatory variable. As far as the US-specific pass-through coefficients are concerned, results are mixed : the assumption of a significant differentiated pass-through behaviour towards the USA is accepted by the data in 40 % of Italian exporting industries but less than 20% of Japanese ones (Line 1 in table 8). Whatever the exporting country then, this hypothesis is accepted in a limited number of industries. One possible explanation of those mixed results could be that the USA are not necessarily the main partner of each exporting industry, even if they are on average. In that case, the US-specific pass-through coefficients could be significant only in those industries where they are effectively dominant. To ask this question explicitly, the estimations summarized under the name "Large Partners" in the tables include exchange rate series computed by keeping only data relative to the 5 largest partners (in terms of exported value) for each industry-specific sample. The hypothesis that is explicitly tested is thus that of a differentiated pass-through strategy towards the main partners. As in the previous estimations, the general results are not strongly affected by the addition of this explanatory variable. However, here again, the share of significant group-specific pass-through coefficients is limited to around 1 industry out of 3, and the direction of the effect is ambiguous. Both specifications thus suggest the same thing concerning the impact of the destination market size on the pass-through strategy : if one cannot rule out the possibility that exporters take into account the size of the destination market when deciding on their pass-through strategy, this does not lead systematically to a different pass-through coefficient at the sectorial level.

Another theoretical rationalization of differentiated pass-through strategies according to the destination country is suggested by Devereux and Engel [2001] and underscores the role of the nominal volatility on the exporter's optimal pass-through strategy : "a country that has highly volatile monetary policy [...] will experience a high rate of pass-through from exchange rate to imported goods". The effect of volatility is also implicit in Froot and Klemperer [1989]'s model that predict the pass-through to be lower when nominal exchange rate variablity is high and exporters try to maintain local market share. To ask for this kind of effect, two additional estimations have been done. In the first one, a "Euro Zone"-specific exchange rate series is computed [20], to ask for the existence of a specific pass-through coefficient towards those countries. With regards to the considered hypothesis, results concerning German, French and Italian exporters are especially interesting. Indeed, during the period, the future members of the Monetary Union have

20. That contains data relative to Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal and Spain. Greece is not included because its membership to the European Monetary Union was only confirmed in 2001 and its currency during the period is highly volatile.

tried (and succeeded) to stabilize their currencies. If Froot and Klemperer's intuition is true, Italian, German and French exporters should then have an incentive to adapt their pass-through strategy to this stable environment. Results can however be ambiguous because the low volatility is not the only feature of the grouped countries : additional efforts have also been done to promote trade in a competitive environment (thus increasing arbitrage possibilities and limiting firms' capacity to discriminate their markets), to favor "healthy" macroeconomic policies, etc... The last estimation thus tests the opposite part of the hypothesis, the specificity of the pass-through strategy towards the highly volatile countries. This is done by constructing a group-specific exchange rate series including Greece, Hungary, North Korea, Mexico, Poland and Turkey, the six importing countries whose volatility towards the dollar is the highest during the period. Here again, the interpretation can be ambiguous since those countries are also among the less developed.

The comparison of those estimations results with those of the initial specification are less satisfactory than previously. Indeed, results are strongly affected by the addition of the group-specific variables. As far as the Euro zone series is concerned, the sensitivity of the results mainly concerns German and French data : Tables 4 and 5 show that both the Hausman statistic and the sign of the mean pass-through coefficient are more affected by this additional variable than by the three others. An interpretation of this high sensitivity of the results is that German and French estimated coefficients under the homogeneity assumption could be strongly influenced by their behaviours towards their European partners. However, here again, the share of "Euro-specific" pass-through coefficients that are significantly different from 0 as well as their sign is too ambiguous to confirm the existence of a systematic differentiated pass-through strategy.

The sensitivity of the results is maximum when controlling for the effect of highly volatile countries ; in particular, the range of coefficients is much wider than in the benchmark estimation (see comparison of lines 4 and 1 in table 7). The presence of those importing countries could thus explain the weakness of the estimated pass-through observed in the previous subsection. Fortunately, this does not change the main qualitative results : the German distribution is still more concentrated around zero than the five others, Italian exporters still appear to have a higher tendency to stabilize local prices and coefficients are still highly heterogeneous from one industry to another. Moreover, if results are sensitive to the presence of those group-specific variables, the existence of a differentiated pass-through strategy towards these specific countries is not confirmed by the data : the specific pass-through coefficients relative to the highly volatile countries are significant in only 40% of the considered industries.

To summarize, the results of this sensitivity analysis lead to the following conclusions. At the sectorial level, nor the size of the destination market, neither the exchange rate volatility appear to have a strong enough influence on pass-through decisions to lead in a recurrent way to a significant differentiated pass-through strategy. One possible explanation of this is that the influence of those macroeconomic features must differ from an industry to another according to structural specificities, as for instance the exposure to currency movements, the market power of firms, etc... Finally, those results confirm the main intuition that emerged from the previous subsection, that the determinants of pass-through strategies should be rather sought at the microeconomic level.

4 Conclusion

Using disaggregated data on a large array of industries and several exporting countries, the previous systematic estimations give us a rich information about the effect of exchange rate fluctuations on international trade prices. After controlling for the possible effect of supply shocks and importer-specific effects, export prices appear sensitive to exchange rate movements in around 40 to 50% of the industries, whatever the exporting country. Conflicting with the standard hypothesis of a full pass-through of exchange movements into import prices, this share is large enough to have an impact at the global level, thus confirming the importance of this issue for macroeconomists. When comparing exporting countries, slight differences appear in the propensity of each country to price-to-market. However, the high heterogeneity of coefficients both between industries (whatever the considered degree of aggregation) and between exporters in a given industry prevents us to interpret it as a macroeconomic phenomenon. The disparity of results rather suggests that structural features such as the firm competitive environment have a large influence on the pass-through decisions. This intuition is confirmed by the robustness checks that fail to put in evidence systematic differences in pass-through strategies towards the USA or the Euro zone, nor towards the main partners or the highly volatile countries.

From a theoretical point-of-view, those results cast doubts on the realism of models that explain the incomplete pass-through phenomenon by the macroeconomic environment in the exporting or the importing country. If those determinants can still influence pass-through strategies, the problem must be tackled in a well microfunded framework that explicitly describes the exporters behaviours[21]. Ideally, such a model should take into account

21. As for instance, in Corsetti and Dedola [2003] that explain strategies of incomplete

the heterogeneity of pass-through strategies, by contemplating various market structures potentially explaining the highest incentives to price-to-market in some given industries (where the competitive pressure pushes to stabilize local prices). Last, the standard theoretical framework inherited from the pricing-to-market literature should also be completed to authorize richer pass-through behaviours such as local currency price amplifications as observed in various industries.

pass-through by introducing local distribution costs.

A1. Derivation of the estimated equation from optimal prices

The optimal price fixed by a representative firm in a given national industry for its sales in the country j during the period t can be written as:

$$P_t^j = \mu_t^j Cm_t$$

with μ_t^j the optimal mark-up that depends on the perceived elasticity of demand η_t^j:

$$\begin{aligned} \mu_t^j &= \tfrac{\eta_t^j}{\eta_t^j - 1} \\ \eta_t^j &= -\tfrac{\partial D_t^j / D_t^j}{\partial (P_t^j S_t^j)/(P_t^j S_t^j)} \quad = \quad \eta^j(P_t^j S_t^j, Z_t^j) \end{aligned}$$

By log-linearizing this equation, one obtains:

$$p_t^j = \ln \eta^j(P_t^j S_t^j, Z_t^j) - \ln(\eta^j(P_t^j S_t^j, Z_t^j) - 1) + cm_t$$

where lower cases indicate the logarithms of the variables in levels. Equation (2) is then obtained from a Taylor approximation of $\ln\left(\frac{\eta^j(P_t^j S_t^j, Z_t^j)}{\eta^j(P_t^j S_t^j, Z_t^j)-1}\right)$ around a given point where $P_t^j S_t^j = P^j S^j$ and $Z_t^j = Z^j$:

$$p_t^j \quad = \quad cm_t + \ln \frac{\eta_0^j}{\eta_0^j - 1} + \left(1 - \frac{\eta_0^j}{\eta_0^j - 1}\right) \varepsilon_{P^j S^j}^{\eta^j} (p_t^j + s_t^j) + \left(1 - \frac{\eta_0^j}{\eta_0^j - 1}\right) \varepsilon_{Z^j}^{\eta^j} \; z_t^j$$

where:

$$\begin{aligned} \eta_0^j &= \eta^j(P^j S^j, Z^j) \\ \varepsilon_{P^j S^j}^{\eta^j} &= \left. \frac{\partial \ln[\eta(P_t^j S_t^j, Z_t^j)]}{\partial \ln(P_t^j S_t^j)} \right|_{P_t^j S_t^j = P^j S^j,\; Z_t^j = Z^j} \\ \varepsilon_{Z^j}^{\eta^j} &= \left. \frac{\partial \ln[\eta(P_t^j S_t^j, Z_t^j)]}{\partial \ln Z_t^j} \right|_{P_t^j S_t^j = P^j S^j,\; Z_t^j = Z^j} \end{aligned}$$

Rearranging the terms of this equation to isolate the export price leads immediately to equation (2):

$$\begin{aligned} p_t^j &= \frac{\eta_0^j - 1}{\eta_0^j - 1 + \varepsilon_{P^j S^j}^{\eta^j}} \left(cm_t + \ln \frac{\eta_0^j}{\eta_0^j - 1} \right) - \frac{\varepsilon_{P^j S^j}^{\eta^j}}{\eta_0^j - 1 + \varepsilon_{P^j S^j}^{\eta_0^j}} s_t^j - \frac{\varepsilon_{Z^j}^{\eta^j}}{\eta_0^j - 1 + \varepsilon_{P^j S^j}^{\eta^j}} z_t^j \\ \Rightarrow p_t^j &= (1+\beta^j) cm_t + (1+\beta^j) \ln \frac{\eta_0^j}{\eta_0^j - 1} + \beta^j s_t^j + \frac{\varepsilon_{Z^j}^{\eta^j}}{\varepsilon_{P^j S^j}^{\eta^j}} \beta^j z_t^j \end{aligned}$$

where $\beta^j \equiv -\frac{\varepsilon_{P^j S^j}^{\eta^j}}{\eta_0^j - 1 + \varepsilon_{P^j S^j}^{\eta^j}}$

A2. Details on the econometric method

The panel equation estimated in this paper is a random effects model of the following form :

$$
\begin{aligned}
p_t^j &= \alpha_t + \beta s_t^j + \gamma^j + \varepsilon_t^j \quad (1) \\
\Leftrightarrow p_t^j &= X_t^j b + u_t^j
\end{aligned}
$$

with $u_t^j = \gamma^j + \varepsilon_t^j$. The estimation of such a model is very standard, except from the fact that the panel is generally seen as balanced whereas it is not the case here : the number of periods varies from an individual to another since every country does not necessarily import goods in every year.

Write the model for all time periods as :

$$p^j = X^j b + u^j$$

where

- p^j is the $(T^j$,1$)$ matrix of the dependent variable, associated with the individual j, present in the panel during T^j years,
- X^j is a $(T^j$,$p)$ matrix of p explanatory variables (in the present case, the exchange rate and T time effects, with $T = \max_j\{T^j\}$),
- b is the $(p;1)$ corresponding matrix of coefficients, supposed identical for all j in the estimation,
- and u^j is the $(T^j$,1$)$ matrix of errors that can be decomposed into two parts, $u^j = \gamma^j e_{T^j} + \varepsilon^j$ [22]

Assume that :

1. $E(\varepsilon^j) = 0$, $E(\varepsilon^j \varepsilon^{j\prime} | X^j, \gamma^j) = \sigma_\varepsilon^2 I_{T^j}$,
2. $E(\gamma^j | X_t^j) = E(\gamma^j) = 0$ (with the assumption $E(\gamma^j) = 0$ being without loss of generality, provided an intercept is included in X_t^j, as is the case here) and $E(\gamma^{j\,2} | X^j) = \sigma_\gamma^2$,

Under these assumptions, one obtains directly the unconditional variance matrix of u^j:

$$\Sigma^j = E(u^j u^{j\prime}) = \sigma_\varepsilon^2 I_{T^j} + \sigma_\gamma^2 e_{T^j} e_{T^j}'$$

This matrix is characterized by a correlation between the composite errors u_t^j and u_s^j, that however does not depend on the difference between t and s: $Corr(u_t^j, u_s^j) = \frac{\sigma_\gamma^2}{\sigma_\gamma^2 + \sigma_\varepsilon^2}$, $\forall t \neq s$. Because this correlation is not equal to 0, the

22. In the following, we will use the following notations : e_n is a $(n$,1$)$ vector of 1, I_n is the identity matrix of size n.

Ordinary Least Squares (OLS) estimation is not efficient [23], whereas Generalized Least Squares (GLS) are consistent and efficient, under the previous assumptions. Since the exact matrices of variance-covariance Σ^j are unknown however, the method I use is the Feasible Generalized Least Squares, leading to the following estimator :

$$\hat{b}_{FGLS} = \left(\sum_{j=1}^{N} X^{j\prime} \hat{\Sigma^j}^{-1} X^j \right)^{-1} \left(\sum_{j=1}^{N} X^{j\prime} \hat{\Sigma^j}^{-1} p^j \right)$$

In order to implement the FGLS procedure, $\hat{\sigma}^2_\gamma$ and $\hat{\sigma}^2_\varepsilon$ are estimated using the between and the within transformations of the model. The between transformation is obtained by averaging equation (1) over $t = 1,...,T^j$ to get the cross section equation :

$$\bar{p^j} = \bar{X^j} b + u^j_B \quad (2)$$

where $\bar{X^j} = (\bar{X}^j_1,...,\bar{X}^j_p)$ and $\bar{x^j} = \frac{\sum_{t=1}^{T^j} x^j_t}{T_j}$. With the matrix language, one obtains :

$$B^j p^j = B^j X^j b + B^j u^j$$

$$Bp = BXb + Bu$$

with :

$$B = \begin{pmatrix} B^1 & 0 & \dots & 0 \\ 0 & B^2 & & \\ \vdots & & \ddots & \\ 0 & & & B^N \end{pmatrix}$$

$$B^j = \frac{e_{T^j} e'_{T^j}}{T^j}$$

The within specification is a time demeaning transformation :

$$p^j_t - \bar{p^j} = (X^j_t - \bar{X^j})b + u^j_{Wt} \quad (3)$$

that can also be written using a transformation matrix :

$$W^j p^j = W^j X^j b + W^j \varepsilon^j$$

$$Wp = WXb + W\varepsilon$$

23. Indeed, $V(\hat{b}_{OLS}) = (X'X)^{-1} X' V(u) X (X'X)^{-1}$

where

$$W = \begin{pmatrix} W^1 & 0 & \dots & 0 \\ 0 & W^2 & & \\ \vdots & & \ddots & \\ 0 & & & W^N \end{pmatrix}$$

$$W^j = I_{T^j} - B^j$$

Note that both B^j and W^j are idempotent and symmetric.

A2.1. The Within estimator

The OLS estimator of the within equation leads to $\hat{\sigma}_\varepsilon^j$ using the estimated residuals written as

$$\hat{u}_W = Wp - WX\hat{b}_W = M_W W u$$

with

$$M_W = I_{NT} - X(X'WX)^{-1}X'$$

where $NT = \sum_{j=1}^{N} T^j$ is the total size of the panel. Indeed, using these notations, the unconditional variance of the estimated residuals is:

$$\begin{aligned} E(\hat{u}_W'\hat{u}_W) &= E(tr(u'WM_WWu)) \\ &= tr(WM_WWE(u'u)) \end{aligned}$$

(with tr the operator of the trace). Using the expressions of Σ^j for $j = 1,...,N$, one can verify that: $WE(u'u) = \sigma_\varepsilon^2 W$. The previous expression can thus be re-write as:

$$\begin{aligned} E(\hat{u}_W'\hat{u}_W) &= \sigma_\varepsilon^2 tr(M_W W) \\ &= \sigma_\varepsilon^2 (NT - N - p + 1) \end{aligned}$$

One obtains then easily an estimation of the variance of the residual:

$$\hat{\sigma}_\varepsilon^j = \frac{\hat{u}_W'\hat{u}_W}{NT - N - p + 1}$$

A2.2. The Between estimator

The standard way of doing is to estimate the variance of the error of this model that leads to $\hat{\sigma_\gamma^2}$, when combined with $\hat{\sigma_\varepsilon^2}$.

Indeed, the OLS estimated residuals of this model can be written as :

$$\hat{u_B} = Bp - BX\hat{b_B} = M_B B u$$

with

$$M_B = I_{NT} - BX(X'BX)^{-1}X'B$$

The unconditional variance of the residuals is then :

$$\begin{aligned} E(\hat{u}_B'\hat{u}_B) &= E(tr(u'BM_BBu)) \\ &= tr(BM_BBE(u'u)) \end{aligned}$$

In the case of a balanced panel (when $T^j = T \ \forall j$), one can verify that $BE(u'u) = (\sigma_\varepsilon^2 + T\sigma_\gamma^2)B$. In that case, the unconditional variance becomes :

$$\begin{aligned} E(\hat{u}_B'\hat{u}_B) &= (\sigma_\varepsilon^2 + T\sigma_\gamma^2)tr(BM_BB) \\ &= (\sigma_\varepsilon^2 + T\sigma_\gamma^2)(NT - p) \end{aligned}$$

Here however, because the panel is unbalanced, one cannot use directly this expression to derive the estimated variance of the individual effect. A convenient way to deal with the size differences of the individual series is to estimate a transformed version of (2), namely :

$$\sqrt{T^j}\bar{p}^j = \sqrt{T^j}\bar{X}^j b + u_{TB}^j \quad (2')$$

The estimated residuals obtained by the OLS can then be written as :

$$\begin{aligned} \hat{u}_{TB}^j &= \sqrt{T^j}B^j p^j - \sqrt{T^j}B^j X^j \hat{b}_{TB} \\ &= \sqrt{T^j}M_{B^j}B^j u^j \end{aligned}$$

with $M_{B^j} = I_{T^j} - B^j X^j (X^{j\prime}B^j X^j)^{-1}X^{j\prime}B^j$ and the estimated variance of the corresponding residual is :

$$\hat{\sigma}_{TB}^2 = \frac{\hat{u}_{TB}'\hat{u}_{TB}}{\sum_{j=1}^N T^j - p}$$

By analogy with the balanced case, the variance of the individual effects is estimated by :

$$\hat{\sigma}_\gamma^2 = max\left(\frac{\hat{\sigma}_\varepsilon^2 - \hat{\sigma}_{TB}^2}{\bar{T}}; 0\right)$$

with $\bar{T} = \frac{\sum_{j=1}^N T^j}{N}$ the mean number of periods by individual.

A2.3. The FGLS estimator

The previous estimated variances permit to derive an estimate of the parameter θ that is used in the FGLS estimation. Here again, one has to take into account the fact that the panel is unbalanced :

$$\hat{\theta}^{j2} = \frac{\hat{\sigma}_{\varepsilon}^2}{\hat{\sigma}_{\varepsilon}^2 + T^j \hat{\sigma}_{\gamma}^2}$$

The FGLS estimator is thus obtained by applying the OLS to :

$$(W^j + \hat{\theta}^j B^j) p_j = (W^j + \hat{\theta}^j B^j) X^j b + u_{FGLS}^j$$

leading to

$$\hat{b}_{FGLS} = \left(\sum_{j=1}^{N} X^{j\prime} \hat{\Sigma^j}^{-1} X^j \right)^{-1} \left(\sum_{j=1}^{N} X^{j\prime} \hat{\Sigma^j}^{-1} p^j \right)$$

Using results of this estimation, the Hausman test is done to check for the consistency of the FGLS estimator, that is to say for the absence of correlation between the individual effects and the explanatory variables. Since the OLS estimator is consistent even if γ^j and X_t^j are correlated, a statistically significant difference between the OLS and the FGLS estimates is interpreted as evidence against the random effect assumption. The original form of the Hausman test is based on the null of no correlation between the individual effects and the explanatory variables that implies that the GLS estimator is asymptotically efficient. Let $\hat{\delta}_{OLS}$ denote the vector of fixed effects obtained from the OLS regression and $\hat{\delta}_{FGLS}$ the vector of random effects corresponding to the FGLS results. Then, under the null,

$$H = (\hat{\delta}_{OLS} - \hat{\delta}_{FGLS})' [Av\hat{a}r(\hat{\delta}_{OLS}) - Av\hat{a}r(\hat{\delta}_{FGLS})]^{-1} (\hat{\delta}_{OLS} - \hat{\delta}_{FGLS})$$

is distributed asymptotically as χ_{p-1}^2 with :

- $(p-1)$ the number of individual effects,
- $Av\hat{a}r(\hat{\delta}_{OLS}) = \sigma_{\varepsilon}^2 [E(X^{j\prime} B^j X^j]^{-1}/N$
- and $Av\hat{a}r(\hat{\delta}_{FGLS}) = \sigma_{\varepsilon}^2 [E(X^{j\prime}(W^j + \theta^j B^j) X^j]^{-1}/N$.

(see Chapter 10 of Wooldridge [2002] for details).

A3. Biased estimate in the case of serial correlation

When the errors are correlated because of the omission of the lagged price in the estimated equation, pass-through coefficients, whether estimated by an OLS or by a FGLS method, are biased. To see that, take the estimated equation (4) ($p_t^j = \alpha_t + \gamma^j + \beta^j s_t^j + \varepsilon_t^j$) and assume that the residual term is non orthogonal because of price rigidities ($\varepsilon_t^j = \delta p_{t-1}^j + \epsilon_t^j$ with ϵ_t^j $i.i.d.$). When assuming that ε_t^j is orthogonal, one obtains biased estimates of β^j:

$$\begin{aligned}
\hat{\beta^j_{OLS}} &= \beta + \delta \frac{\sum_{i=1}^N \sum_{t=1}^T s_t^j p_{t-1}^j}{\sum_{i=1}^N \sum_{t=1}^T {s_t^j}^2} + \frac{\sum_{i=1}^N \sum_{t=1}^T s_t^j \epsilon_t^j}{\sum_{i=1}^N \sum_{t=1}^T {s_t^j}^2} \\
\hat{\beta^j_{FGLS}} &= \beta + \delta \frac{\sum_{i=1}^N \sum_{t=1}^T s_t^j \hat{\Sigma}^j p_{t-1}^j}{\sum_{i=1}^N \sum_{t=1}^T \hat{\Sigma}^j {s_t^j}^2} + \frac{\sum_{i=1}^N \sum_{t=1}^T s_t^j \hat{\Sigma}^j \epsilon_t^j}{\sum_{i=1}^N \sum_{t=1}^T \hat{\Sigma}^j {s_t^j}^2}
\end{aligned}$$

The expected biased is then of the following form :

$$\begin{aligned}
E(\hat{\beta_{OLS}} - \beta) &= \delta \frac{\sum_{i=1}^N \sum_{t=1}^T s_t^j p_{t-1}^j}{\sum_{i=1}^N \sum_{t=1}^T {s_t^j}^2} \\
E(\hat{\beta_{FGLS}} - \beta) &= \delta \frac{\sum_{i=1}^N \sum_{t=1}^T s_t^j \hat{\Sigma}^j p_{t-1}^j}{\sum_{i=1}^N \sum_{t=1}^T \hat{\Sigma}^j {s_t^j}^2}
\end{aligned}$$

As explained in Wooldridge [2002], the sign of the bias depends on the covariance of the omitted and the explanatory variables (of the lag of the price and the current exchange rate in our case) :

$$plim \hat{\beta} = \beta + \delta \frac{Cov(s_t^j, y_{t-1}^j)}{Var s_t^j}$$

As the persistence of prices would imply $\delta > 0$, one can see that the bias due to the omission of the lagged price is positive as long as the nominal exchange rate and the lagged price are positively correlated.

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Tab. 1 – *Statistics on the individual exchange rate series (with regards to the dollar)*

Importing country	**Mean annual growth rate**	**Standard error**
Australia	.025	.072
Austria	-.007	.093
Belgium	-.009	.098
Canada	.025	.034
Switzerland	-.014	.103
Czech Republic	.020	.086
Germany	-.007	.091
Denmark	-.010	.094
Spain	.026	.108
Finland	.024	.132
France	-.009	.089
United Kingdom	-.002	.069
Greece	.067	.075
Hungary	.143	.052
Ireland	-.002	.085
Iceland	.024	.065
Italy	.026	.104
Japan	-.006	.100
Korea	.082	.128
Mexico	.146	.194
Netherlands	-.008	.093
Norway	.010	.082
New Zealand	.013	.095
Poland	.354	.581
Portugal	.015	.094
Sweden	.023	.121
Turkey	.535	.208

TAB. 2 – *Averaged statistics on the distributions of the sector-based nominal exchange rate pass-through (500 coefficients considered)*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
Mean estimated standard error	0.030	0.041	0.020	0.026	0.046	0.034
Share of significant coefficients (%)[(a)]	41	52	65	51	52	57
Share of negative significant coefficients (%)	59	53	76	67	69	62
Interquartile range of coefficients[(b)]	-0.29;0.19	-0.25;0.23	-0.18;0.08	-0.24;0.10	-0.59;0.13	-0.31;0.22

(a) Fraction of estimated coefficients that are significantly different from 0 at the 5% level
(b) Range of the estimated coefficients ignoring the tails (10% of the coefficients ignored)

TAB. 3 – *Averaged statistics on the distributions of the sector-based real exchange rate pass-through*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
Mean estimated standard error	0.049	0.078	0.067	0.067	0.088	0.056
Share of significant coefficients (%)	40	52	65	54	52	56
Share of negative significant coefficients (%)	52	53	34	45	74	46
Interquartile range of coefficients	-0.63;0.11	-0.69;0.84	-0.36;0.76	-0.55;0.78	-1.07;0.07	-0.60;0.68

Tab. 4 – *Effect of the addition of a group specific exchange rate series on the chosen estimated model (OLS or FGLS) : share of industries in which the chosen model is the same as in the benchmark estimation (%)*[(a)]

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
USA	.	88.0	88.9	88.1	82.9	88.7
Euro Zone	86.3	85.7	47.0	48.2	80.3	82.2
High Volatility	86.3	84.5	84.3	85.7	80.5	78.7
Large Partners	88.5	88.3	85.2	86.4	84.7	85.0

(a) All estimations are based using an equation of the following form :

$$\ln P_t^j = \alpha_t + \gamma^j + \beta \ln S_t^j + \beta^j D^j \ln S_t^j + \varepsilon_t^j$$

where D^j is a dummy variable computed so that it equals 1 when the observation concerns an importing country that is included in the studied group. The composition of groups is the following :

1. the United States,
2. Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal and Spain,
3. Greece, Hungary, North Korea, Mexico, Poland and Turkey,
4. 5 largest partners (in terms of exported value) of each industry- and exporter-specific sample.

Tab. 5 – *Effect of the additional explanatory variable on the sign of common pass-through coefficients : share of industries in which the sign of the estimated coefficient is the same as in the benchmark estimation (%)*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
USA	.	88.8	90.0	87.9	86.7	88.7
Euro Zone	80.9	79.0	59.6	59.0	85.7	81.0
High Volatility	73.6	70.8	63.3	67.0	82.8	70.1
Large Partners	86.9	86.7	88.8	84.3	88.3	82.1

TAB. 6 – *Share of common coefficients (β) that are significantly different from zero in each specification (%)*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
Benchmark	40.6	52.4	65.2	51.0	52.0	56.8
USA	.	47.2	65.2	48.5	45.0	54.6
Euro Zone	34.2	46.4	63.0	49.8	43.7	49.3
High Volatility	36.3	41.3	38.0	41.0	48.7	44.4
Large Partners	37.0	49.7	63.6	50.2	48.9	54.1

TAB. 7 – *Interquartile ranges* [(a)] *of the common pass-through coefficients (β)*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
Benchmark	-.28;.19	-.25;.22	-.18;.08	-.24;.10	-.59;.13	-.31;.22
USA	.	-.24;.15	-.18;.06	-.24;.07	-.57;.13	-.30;.16
Euro Zone	-.18;.17	-.23;.19	-.19;.08	-.25;.10	-.58;.19	-.35;.19
High Volatility	-.85;.48	-.76;.68	-.51;.50	-.62;.47	-.98;.08	-.71;.50
Large Partners	-.28;.25	-.24;.23	-.19;.06	-.24;.11	-.69;.19	-.34;.21

(a) 90% of the total distribution of coefficients taken into account.

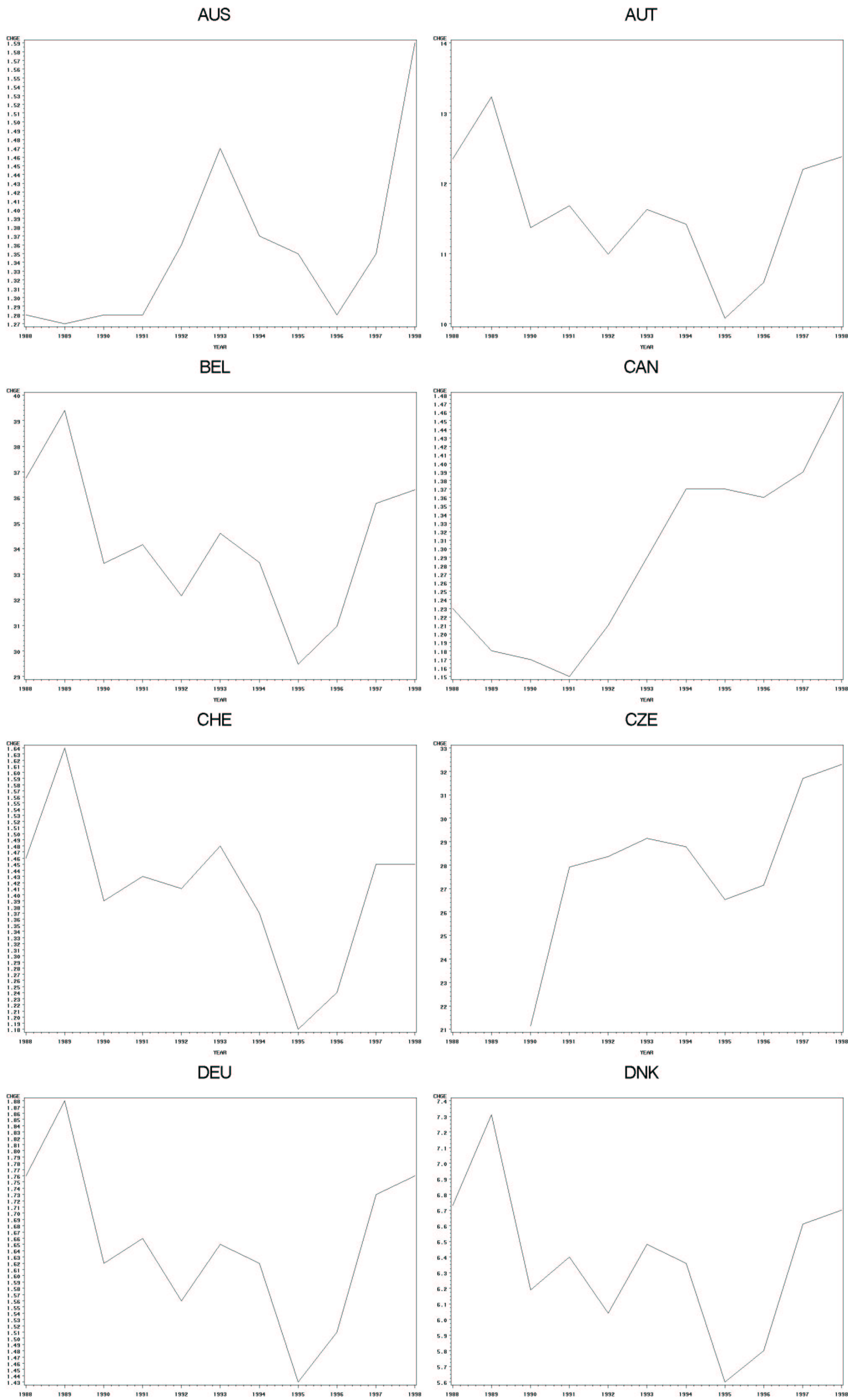
TAB. 8 – *Share of group-specific coefficients (β^j) that are significantly different from zero (%)*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
USA	.	38.0	30.3	37.9	40.5	19.6
Euro Zone	37.0	47.8	41.0	43.2	44.9	43.6
High Volatility	38.7	41.7	39.6	42.8	37.8	44.0
Large Partners	31.3	37.6	30.2	33.7	40.8	42.0

TAB. 9 – *Share of the significant group-specific coefficients (β^j) that are negative (%)*

	USA	**UK**	**DEU**	**FRA**	**ITA**	**JAP**
USA	.	28.4	46.6	63.1	66.1	54.6
Euro Zone	54.4	61.6	47.3	53.2	41.7	48.8
High Volatility	46.9	41.7	51.3	45.0	31.3	53.9
Large Partners	49.0	35.8	49.7	53.0	51.5	46.6

FIG. 1 – *Evolution of the importing countries' exchange rate with regards to the dollar*



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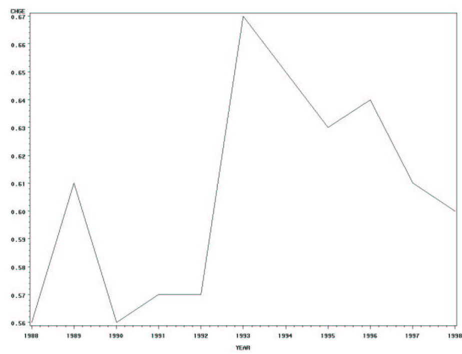

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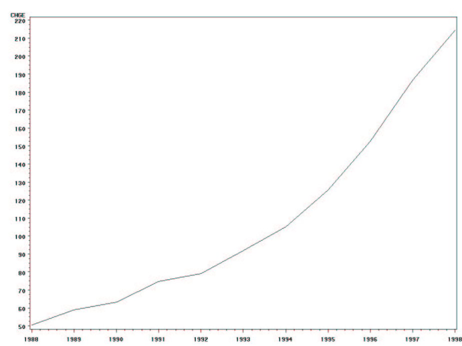

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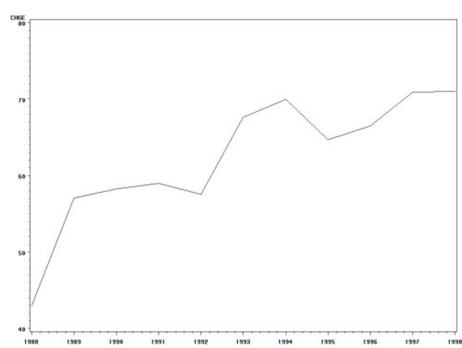

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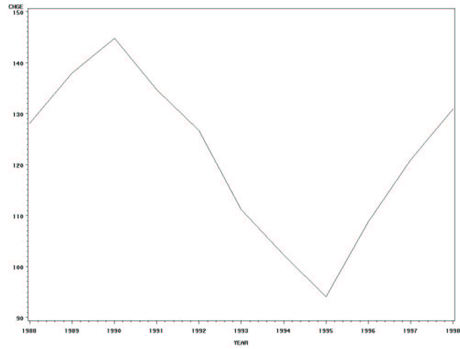

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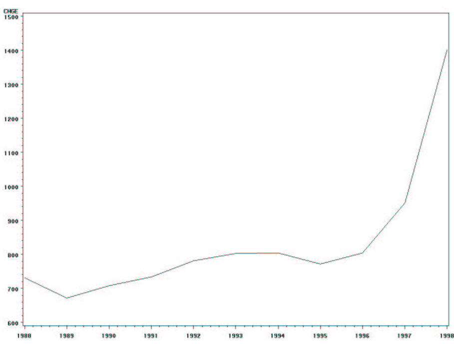

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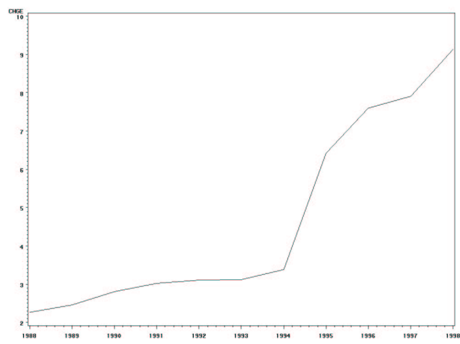

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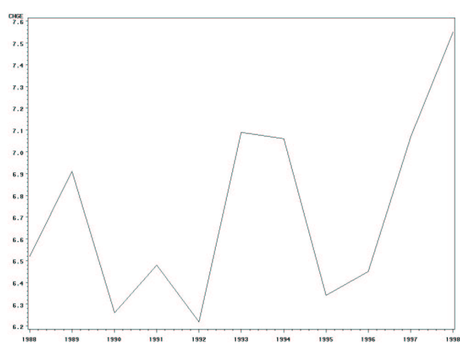

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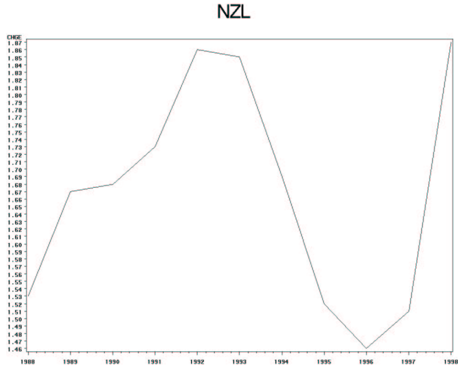

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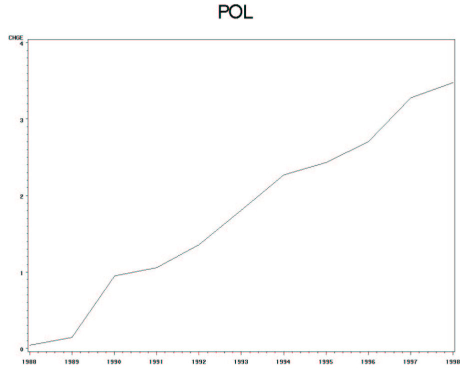

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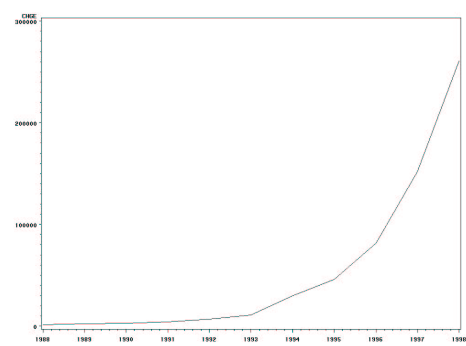
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Fig. 2 – *Estimated distributions of the nominal exchange rate pass-through*