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Import Prices, Variety and the Extensive Margin of Trade

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IMPORT PRICES, VARIETY AND THE EXTENSIVE MARGIN OF TRADE

SUMMARY

Explaining the determinants of relative prices is a recurrent matter of interest in international macroeconomics. Recently, several models have used tools of the New Trade Theory to emphasize the price impact of firms' entry decisions in national markets. In these models, the increased supply of variety attributable to the entry of new firms in a market has a depreciating effect on prices. This sensitivity of national price levels to endogenous entry decisions on national price levels has interesting implications in terms of the long-run determinants of real exchange rates.

This literature however lacks empirical evidence concerning these aggregate price effects. Indeed, intuitions of these models must be tested on aggregate price series but available data are usually constructed using price index formula that do not catch new varieties. In this paper, we use a method, first proposed by Feenstra (1994) and recently applied to US data by Broda & Weinstein (2005), that allows to quantify the effect that newly imported varieties have on import price indices. We apply it to the BACI database that describes bilateral trade flows in more than 5000 groups of products among some 130 countries. This allows us to compare the price impact of changes in the supply of variety in a sample of importing countries. From this cross-country comparison, results in terms of *relative* prices can be gathered.

Our empirical strategy is constructed on a theoretical framework describing import price levels when consumers value variety and the supply of variety for each consumption good is allowed to vary. In this framework, an increase in the share of newly imported varieties in the consumed value of a given good leads to an aggregate price drop, which is all the stronger as varieties of this good are poor substitutes. Such an effect is not caught by price indices constructed on a constant basket of varieties, which leads to a measurement error. To quantify this bias, we apply Broda & Weinstein (2005) empirical strategy that can be decomposed into several steps. First, we estimate the elasticity of substitution between varieties at the good- and country-level, using a panel data method. Then, still at the sectorial level, we measure the price effect of new varieties during the 1994-2003 period for each of the considered importing country. Last, we aggregate these sectorial biases at the country level to obtain a measure of the aggregate price effect of changes in the supply of variety for imported goods. The cross-country comparison of these aggregate biases allows us to infer results in terms of relative prices.

Our results show that changes in the variety supply of imported goods generally have a negative impact on price indices that is not caught by conventional price measures based on a constant basket of varieties. This measurement bias is limited on average, 0.2% a year between 1994 and 2003. However, the results strongly vary across sectors as well as across importing countries. The measurement bias is significantly higher in some emerging countries like India, Indonesia and Brazil whereas it is weaker in most European countries. This suggests that real exchange rates in fast growing countries are more depreciated than measured in standard statistics.

Our interpretation of this result is that fast growing countries are considered by exporters as promising markets, which gives them the incentive to pay the fixed cost to enter these markets. This entry of new firms through exports increases the supply of variety available to local consumers which has a decreasing effect on import prices.

ABSTRACT

This paper studies the aggregate price effect of newly imported varieties and compares it in a sample of countries. The method allows to quantify the measurement bias in import price indices that take as given the basket of imported varieties and neglect the aggregate effect of increased diversity. Applying it to the BACI database describing bilateral trade flows at the world level, we are able to compare the aggregate price impact of the extensive margin of trade among 28 countries. Our results suggest that, in the 1994-2003 period, neglecting newly imported varieties leads to overestimate the import price level by 0.2% a year, on average. The magnitude of this effect however strongly varies across countries, this overestimation being especially strong in some emerging countries like India, Indonesia or Brazil.

JEL Classification: F10, F12, F41
Keywords: extensive margin, import price indices, real exchange rate determinants, panel data

PRIX À L'IMPORTATION, OFFRE DE VARIÉTÉ ET MARGE EXTENSIVE DU COMMERCE

RÉSUMÉ

La question des déterminants des prix relatifs est récurrente en macroéconomie internationale. Plusieurs modèles ont récemment utilisé les outils des nouvelles théories du commerce pour mettre en évidence l'impact, sur les prix, des décisions d'entrée des firmes sur les marchés nationaux. Dans ces modèles, l'augmentation de l'offre de variétés attribuable aux décisions d'entrée des firmes sur un marché conduit à une baisse des prix. Une telle sensibilité du niveau des prix aux choix d'entrée sur un marché a des implications intéressantes en ce qui concerne les déterminants du taux de change réel de long terme.

Les preuves empiriques de l'existence de tels effets manquent cependant. Les intuitions de ces modèles ne peuvent en effet être testées que sur des indices de prix agrégés mais les données disponibles sont généralement construites en utilisant des formules d'indices ne permettant pas d'identifier l'impact des nouvelles variétés. Dans cet article, nous utilisons une méthode proposée par Feenstra (1994) et appliquée récemment sur données américaines par Broda & Weinstein (2005), qui permet de quantifier l'impact sur l'indice de prix à l'importation des variétés nouvellement importées d'un bien. Cette méthode est appliquée à la base de données BACI qui décrit le commerce bilatéral d'environ 130 pays dans plus de 5000 groupes de produits. On peut ainsi comparer l'impact sur les prix des variations endogènes de l'offre de variétés disponibles dans un échantillon de 28 pays importateurs. Cette comparaison internationale permet alors d'inférer l'impact des ces variations sur les prix *relatifs*.

Notre stratégie empirique utilise un cadre analytique décrivant le niveau des prix à l'importation lorsque les consommateurs ont une préférence pour la variété et que l'offre de variétés des différents biens consommés varie au cours du temps. Sous ces hypothèses, une hausse de la part des nouvelles variétés dans la consommation en valeur d'un bien particulier conduit à une baisse de l'indice de prix agrégé, qui est d'autant plus forte que les variétés du bien considéré sont peu substituables. Un tel ajustement des prix n'est pas pris en compte dans les indices de prix construits sur la base d'un panier de variétés constant, qui sont donc des mesures biaisées de l'indice de prix. Nous quantifions ce biais en utilisant la stratégie empirique de Broda & Weinstein (2005) qui peut être décomposée en plusieurs étapes. D'abord, en utilisant une méthode de panel appliquée à chaque sous-échantillon sectoriel relatifs aux importations d'un pays particulier, nous estimons l'élasticité de substitution entre variétés d'un bien particulier. Ensuite, nous mesurons, toujours au niveau sectoriel, l'effet sur les prix de l'apparition de nouvelles variétés au cours de la période 1994-2003. Enfin, ces biais sectoriels sont agrégés pour obtenir une mesure de l'impact des changements de l'offre de variétés sur l'indice de prix à l'importation de chacun des pays considérés. La comparaison internationale de ces biais agrégés nous permet alors d'inférer les résultats en termes de prix relatifs.

Les résultats montrent que les variations de l'offre de variétés importées ont généralement un impact négatif sur les indices de prix, effet qui n'est pas pris en compte par les mesures conventionnelles des indices de prix qui supposent un panier de variétés constant. Ce biais de mesure est relativement faible en moyenne, 0.2% par an entre 1994 et 2003. Cependant,

les résultats varient fortement entre secteurs ainsi que d'un pays importateur à l'autre. Ainsi, le biais de mesure est nettement plus important dans des pays émergents tels que l'Inde, l'Indonésie ou le Brésil tandis qu'il est plus faible dans la plupart des pays européens. Ces écarts entre pays suggèrent que les taux de change réel des pays émergents sont susceptibles d'être plus dépréciés que ce qui est mesuré par les statistiques officielles.
Notre interprétation de ce résultat est que les pays émergents représentent des potentiels de marchés prometteurs, ce qui incite les exportateurs à payer le coût fixe d'entrée sur ces marchés. Cette entrée de nouvelles firmes augmente l'offre de variétés à laquelle les consommateurs locaux ont accès, ce qui réduit l'indice de prix à l'importation.

RÉSUMÉ COURT

Cet article étudie l'effet, sur les prix agrégés, de l'arrivée de nouvelles variétés sur les marchés à l'importation et compare l'ampleur de ces ajustements au sein d'un échantillon de pays. La méthode utilisée permet de quantifier le biais de mesure des indices de prix à l'importation qui considèrent que le panier de variétés importées est constant et négligent l'effet agrégé d'une augmentation de la diversité. En l'appliquant à la base de données BACI qui décrit le commerce mondial en bilatéral, nous sommes capables de comparer l'effet de la marge extensive du commerce sur les prix agrégés de 28 pays. Les résultats suggèrent qu'entre 1994 et 2003, le fait de négliger l'arrivée de nouvelles variétés sur les marchés à l'importation a conduit à surestimer l'indice de prix à l'importation de 0.2% par an en moyenne. L'ampleur de cet effet varie cependant fortement entre pays, cette surestimation étant particulièrement prononcée dans des pays émergents tels que l'Inde, l'Indonésie ou le Brésil.

Classification *JEL* : F10, F12, F41
Mots-clé : marge extensive, indices de prix à l'importation, déterminants du taux de change réel, données de panel

Import Prices, Variety and the Extensive Margin of Trade[1]

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1 Introduction

Explaining the determinants of relative prices is a recurrent matter of interest in international macroeconomics. Recently, several models have used tools of the New Trade Theory[4] to study the price impact of firms' entry decisions in national markets.[5] In this paper, we use a method first proposed by Feenstra (1994) that allows to quantify the measurement error attributable to the omission of new varieties in the calculation of price indices. Comparing results across countries makes it possible to measure the sensitivity of relative import prices to endogenous changes in the variety available in each market.
New trade models of the Krugman (1980) type and heterogeneous firms models *à la* Mélitz (2003) have emphasized the role of firms' entry decisions in national markets in explaining aggregate trade flows. In this theoretical framework, trade growth can be decomposed into an "intensive" component attributable to changes in incumbent firms' volume of exports, and an "extensive" component due to the entry/exit of new/disappearing firms in the export market. These models also have implications in terms of prices that may be of interest to macroeconomists. On the one hand, the entry of new firms in the export market might have pro-competitive effects, as discussed by Mélitz & Ottaviano (2005) and Chen, Imbs & Scott (2006). The following aggregate price decrease can be called an "intensive" price effect as it comes from the adjustment of incumbent firms' price strategies. On the other hand, firms' entry decisions may also decrease aggregate price indices through "extensive" effects due to the self-selection of firms in the market (as in Ghironi & Mélitz (2005)) or because of agglomeration effects (as in Corsetti, Martin & Pesenti, 2005). Decomposing price adjustment mechanisms into these "intensive" and "extensive" effects could help explaining one of the main "puzzle" in international macroeconomics, the so-called "PPP Puzzle" that asks for the determinants of long-run deviations of relative prices from their PPP level.[6]

[1]Many thanks to Robert Feenstra for letting us consult his estimation programs. We are very grateful to Lionel Fontagné, Agnès Bénassy-Quéré and participants to the workshop on Exchange rate behaviours organized by the Centro de Estudios Andaluces of Sevilla, for helpful comments.
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[4]See Combes, Mayer & Thisse (2006) for a survey.
[5]See Corsetti, Martin & Pesenti (2005), Ghironi & Mélitz (2005), MacDonald & Ricci (2004) and Méjean (2006).
[6]See Obstfeld & Rogoff (2000).

Whereas several recent models study the theoretical price impact of endogenous location decisions, empirical evidence is still lacking. The intensive price effect has recently been studied by Chen et al. (2006) using disaggregated price data. They find evidence of a competitive effect of trade openness leading to falling markups and prices. However, the extensive price effect has to be tested on suitable aggregate price indices. Indeed, Ghironi & Mélitz (2005) and Corsetti et al. (2005) argue that identifying these effects in conventional price series is tricky because the way these indices are usually constructed does not allow to account for the appearance of new varieties in a market. Indeed, statistical methods for measuring aggregate prices aim at identifying true price variations by neutralizing composition effects, notably those attributable to new or disappearing goods. While they allow to accurately identify individual price variations, these methods omit the effect that endogenous changes in the variety of goods have on aggregate prices. Such a measurement bias can be misleading when estimating import demand elasticities - as argued by Krugman (1991) or Feenstra (1991) - when studying real exchange rate determinants[7] or when trying to measure the impact of increased variety on aggregate welfare (Broda & Weinstein, 2005). One solution to this problem is to incorporate in the right-hand side of estimated equations a control for changes in the supply of variety, as done by Helkie & Hooper (1988) who include a measure of foreign's capital stock as a proxy for the supply of new products. On the other hand, Feenstra, Heston, Timmer & Deng (2004) propose to use a PPP index constructed from output prices to capture changes in countries' production over time. A third solution for correcting this bias, and the one we are using in this paper, has been proposed by Feenstra (1994) and consists in incorporating the new product varieties directly into price index measures.

Feenstra's method uses bilateral trade data to quantify the bias in measured price indices due to the omission of changes in the variety supply of differentiated goods. This empirical strategy catches aggregate price variations attributable to the appearance or disappearance of bilateral trade relations between two dates, taking into account substitution effects as well as the impact working through consumers' taste-for-variety. This method has recently been applied to US data by Broda & Weinstein (2005). In this paper, the authors estimate that "*over the last fifty years, if one adjusts for new varieties, [US] import prices have been falling 1.2 percent faster than one would surmise from official statistics*". Omitting changes in variety supply may thus induce a substantial bias in price indices, which is not surprising since the extensive margin has been shown to explain an important share of trade growth.[8] Moreover, Broda & Weinstein (2004)'s simulations suggest that the size of the variety effect measured in US data could be a lower bound. It should notably be stronger in those countries that recently got involved in a liberalization process (as China and the former Soviet Union). Such cross-country gaps in the magnitude of the price effect of endogoneous variety supplies are interesting because they should lead to a measurement bias of *relative* price levels (that is terms-of-trade and real exchange rates). Studying the price impact of endogenous location decisions for various importing countries is thus of an interest as it could help understanding

[7] See Corsetti et al. (2005), Ghironi & Mélitz (2005), Méjean (2006).

[8] See evidences obtained from sectorial data in Kehoe & Ruhl (2003) or Hummels & Klenow (2002). The extensive margin is also studied using firm-level data by Eaton, Kortum & Kramarz (2004) and Koenig-Soubeyran (2005).

the determinants of relative price levels.

This is what is done in this paper that quantifies the price measurement error due to omitted new varieties in imported markets of 28 countries. To this aim, we apply Feenstra (1994)'s methodology to an exhaustive database covering bilateral trade flows among more than 130 countries in 5,000 sectors, and we compare the results in the geographical dimension. The data are from BACI, a database developed at CEPII that contains informations about bilateral trade flows at the 6-digit sectorial level over the 1994-2003 period. Assuming goods to be differentiated by their country of origin, we first estimate the elasticity of substitution between varieties at the good and importing country level. We are then able to measure sectorial price effects of changes in the variety available in each importing country. Aggregating results let us with an estimation of the measurement bias in the import price index of 28 large importing countries, attributable to the appearance/disappearance of varieties.

The results suggest that changes in the variety of imported goods led to a decrease in price indices that is not caught by conventional price indices based on a constant basket of varieties. This measurement bias is small on average, 0.2% a year between 1994 and 2003. However, the results strongly vary across sectors as well as across the considered importing countries. Indeed, the measurement bias is significantly higher in some emerging countries like India, Indonesia and Brazil whereas it is weaker in most European countries. This suggests that real exchange rates in fast growing countries are even weaker than measured in standard statistics.

The rest of the paper is organized as follows. Section 2 presents the analytical arguments in favour of a measurement bias in price index series that do not take into account the endogeneity of variety supplies. Section 3 details the empirical strategy used to measure the price effect of these changes. In section 4, we present the data as well as some statistics concerning the extent of the extensive margin in this database. Section 5 is devoted to the results. Section 6 concludes.

2 Endogenous variety supply and import price index

2.1 Conventional measures of import prices

Usually, price indices for international trade are constructed using a method that allows to track price trends for a consistent basket of goods.[9] To disentangle "pure" price changes from composition effects, statistical institutes thus follow the price of a limited number of items and aggregate them using price index formula with alternative weighting schemes (generally Laspeyres, Paasche or Fisher). Each time an item disappears, its price is replaced in the index by that of another item that must be as substitutable as possible. In case a good is replaced by a good with substantive different features, the substitution is made by chain linking in order

[9] See for instance the description of US practices on the BLS website, `www.bls.gov/mxp/` and the equivalent in France, `http://www.insee.fr/fr/indicateur/indic_conj/donnees/ivucex_m.pdf`.

the serie not to exhibit a break.[10] Of course, statistical institutes account for the fact that the basket of traded goods changes over time. The sample of products as well as the weighting coefficients are thus steadily revised, at a frequency that strongly varies across countries. In the United States for instance, the sampled products incorporated in the import and export price indices are priced for approximately 5 years until they are replaced.
Whereas such a way-of-doing allows to measure individual price changes in an accurate way, this is far from the best method to investigate the price impact that the appearance of new varieties have on price levels. Indeed, in New Trade Theories, firms entering a market affect price levels through pro-competitive (intensive) effects, as well as through aggregate (i.e. extensive) channels. Whereas the former can be caught by the previously discussed measure of import prices, the latter cannot. As an example, consider Ghironi & Melitz's model which is featured by this kind of "extensive" price effects. In their framework, firms' entry decisions in export markets push import price indices downwards because entering firms are relatively more productive (i.e. more price-competitive) than evicted ones. This sorting mechanism is not caught by price indices if the evicted variety was not in the sample of items the statistical institute follows and if this sample is not revised the year the new firm enters the market. In the same way, all welfare gains linked to the "taste-for-variety" property of New Trade models are ignored by price indices constructed on a constant consumption basket.
In the following, an analytical form of this extensive price effect is derived that will be at the root of the empirical analysis.

2.2 The exact price index with CES preferences

To derive the bias linked to the omission of new varieties in measured price indices, let us consider a country j importing N differentiated goods indexed by k.[11] Consistently with the standard assumption of new trade models, the utility function is assumed to be of the following form:

$$U_j(M_{1jt}, ..., M_{Njt}) = \sum_{k \in K_j} \left(b_{kjt}^{1/\gamma} M_{kjt}^{\frac{\gamma-1}{\gamma}} \right)^{\frac{\gamma}{\gamma-1}} \tag{1}$$

with :

- $K_j \subset \{1, ..., N\}$ the set of imported goods available in country j, which is assumed to be constant over time,
- $b_{kjt} > 0$ a taste parameter for good k,

[10]The chain linking method consists in ignoring the new good in the first pricing period it is introduced and to bring it into the market basket from the second period using the chain approach. Marshall (1887) is credited with suggesting this.

[11]In the following, country j is assumed to consume only imported goods. This restriction is dictated by the type of data used in the empirical part. Indeed, with trade data, we are unable to incorporate domestic products in the analysis. However, the demonstration would be nearly the same if the representative household was assumed to consume a perfectly substitutable domestic good (as in Broda & Weinstein (2005)) or domestically produced varieties of each aggregate k.

- $\gamma > 1$ the elasticity of substitution between imported goods.

In this equation, M_{kjt} measures the sub-utility derived from the consumption of an imported good k at time t, which is written as a non-symmetrical CES function over varieties of this good :

$$M_{kjt} = \sum_{i \in C_{kjt}} \left(d_{kijt}^{1/\sigma_k} m_{kijt}^{\frac{\sigma_k - 1}{\sigma_k}} \right)^{\frac{\sigma_k}{\sigma_k - 1}} \tag{2}$$

where:

- i is a specific variety of good k, identified in the data by the country of origin of the good,[12]
- $C_{kjt} \subset \{1, ..., V_{kjt}\}$ is the set of suppliers (and V_{kjt} the number of suppliers), which varies over time if new varieties start being imported or former ones are no more bought,
- d_{kijt} is a parameter describing the consumer's taste for (or the quality of) the differentiated variety i,
- and $\sigma_k > 1$ is the elasticity of substitution between varieties of this good.

The program of the representative consumer from country j consists in maximizing equation (1) under her budget constraint:

$$\sum_{k \in K_j} \sum_{i \in C_{kjt}} p_{kijt} m_{kijt} \leq E_{jt} \tag{3}$$

where p_{kijt} is the price of variety i at time t and E_{jt} her total consumption expenditure. Solving this leads to the consumed quantity of each variety:

$$m_{kijt} = \left(\frac{p_{kijt}}{P_{kjt}} \right)^{-\sigma} \left(\frac{P_{kjt}}{\phi_{jt}} \right)^{-\gamma} d_{kijt} b_{kjt} \frac{E_{jt}}{\phi_{jt}} \tag{4}$$

where P_{kjt} and ϕ_{jt} are the sectorial and aggregate price indices that minimize the consumer's unit-cost function:

$$P_{kjt} = \left[\sum_{i \in C_{kjt}} d_{kijt} p_{kijt}^{1-\sigma_k} \right]^{\frac{1}{1-\sigma_k}}$$

$$\phi_{jt} = \left[\sum_{k \in K_j} b_{kjt} P_{kjt}^{1-\gamma} \right]^{\frac{1}{1-\gamma}}$$

[12] As explained in the following, Feenstra (1994) and Broda & Weinstein (2005) distinguish varieties together by their country of origin: each exporting country i is assumed to produce a specific (representative) variety. Doing this assumption allows to use the cross-country dimension of bilateral trade data to identify trade in varieties. Moreover, results are not biased if there are more than one variety produced in a given country. See footnote 15.

In the optimum, the aggregate price level thus depends on the price of each consumed variety of imported goods ($\mathbf{p}_{kjt} = \{p_{kijt}\}$), the sets of taste/quality parameters ($\mathbf{b}_{jt} = \{b_{kjt}\}$ and $\mathbf{d}_{kjt} = \{d_{kijt}\}$), the number of imported goods (K_j) and the variety in each product market ($\mathbf{C}_{jt} = \{C_{kjt}\}$).
Using Diewert (1976)'s definition of an exact price index[13] and its extension to the CES case by Sato (1976) and Vartia (1976) and assuming that the taste parameters are constant whereas the set of varieties can change over time (while still overlapping), Feenstra (1994) shows that the exact price index for an individual good k is:

$$P_{kjt}(\mathbf{p}_{kjt}, \mathbf{p}_{kjt-1}, \mathbf{x}_{kjt}, \mathbf{x}_{kjt-1}, C_{kj}) = \left[\prod_{i \in C_{kj}} \left(\frac{p_{kijt}}{p_{kijt-1}} \right)^{w_{kijt}} \right] \left(\frac{\lambda_{kjt}}{\lambda_{kjt-1}} \right)^{\frac{1}{\sigma_k - 1}} \tag{5}$$

with:

- $\mathbf{x}_{kjt}$ the cost-minimizing consumption bundles,
- C_{kj} the set of varieties available in both periods ($C_{kj} = C_{kjt} \cap C_{kjt-1}$),
- $\{w_{kijt}\}$ "ideal log-change weight" (normalized version of the logarithmic means) constructed from each period's cost shares:

$$s_{kijt} = \frac{p_{kijt} x_{kijt}}{\sum_{i \in C_{kjt}} p_{kijt} x_{kijt}}, \quad w_{kijt} = \frac{\frac{s_{kijt} - s_{kijt-1}}{\ln s_{kijt} - \ln s_{kijt-1}}}{\sum_{i \in C_{kj}} \left(\frac{s_{kijt} - s_{kijt-1}}{\ln s_{kijt} - \ln s_{kijt-1}} \right)}$$

- and λ_{kjt} the fraction of period t's expenditures spent on varieties that are available in both periods:

$$\lambda_{kjt} = \frac{\sum_{i \in C_{kj}} p_{kijt} x_{kijt}}{\sum_{i \in C_{kjt}} p_{kijt} x_{kijt}}$$

This price index is convenient as it integers changes in the variety supply in each period (*i.e.* changes in V_{kjt}). Moreover, as λ_{kjt} is constructed from expenditure shares rather than from the number of available varieties, it is not biased by changes in taste parameters.[14]
This formulation clearly identifies the measurement bias occuring when assuming the variety supply of good k to be the same in $t-1$ and t. This bias is measured by the ratio $\lambda_{kjt}/\lambda_{kjt-1}$ raised to the power $1/(1-\sigma_k)$. To understand the mechanisms at work, assume that some new varieties enter country j at time t whereas there is no disappearing variety. Suppose moreover that the entry of new varieties has no effect on prices and quantities. In this case, the "conventional" price index that neglects new varieties overstates the ratio of unit-costs as $\lambda_{kjt} < 1$ and $\lambda_{kjt-1} = 1$. Indeed, the first period price of the new varieties has to be infinite

[13] According to Diewert, an exact price index exactly matches changes in minimum unit-costs.

[14] Indeed, if taste parameters change over time, this price index still represents the ratio of unit-costs with constant taste parameters ($\tilde{d}_{kij}$) lying between normalized versions of d_{kijt-1} and d_{kijt}. See the demonstration in the Appendix of Feenstra (1991).

to rationalize zero demand. Their price thus virtually decreases from infinity to the second period observed price, this price drop being neglected in the conventional price index. As shown by the presence of σ_k in equation (5), this bias is all the more pronounced as varieties are poor substitutes (σ_k low). Indeed, when varieties are highly differentiated, new varieties are very valuable and disappearing varieties very costly from the consumer's viewpoint.

Broda & Weinstein (2005) extend this analysis to the case of multiple imported goods. Assuming that the good-specific taste parameters are constant ($\mathbf{b}_{jt} = \mathbf{b}_j, \forall t$) but letting variety supplies of each good (C_{kjt}) and variety-specific taste parameters ($\mathbf{d}_{kjt}$) vary, they show that the exact import price index (EPI) can be written as follows:

$$EPI(C_{kj}) = \prod_{k \in K_j} P_{kj}^{w_{kjt}(K_j)} = CPI(C_{kj}) \left[\prod_{k \in K_j} \left(\frac{\lambda_{kjt}}{\lambda_{kjt-1}} \right)^{\frac{w_{kjt}(K_j)}{\sigma_k - 1}} \right] \tag{6}$$

where $w_{kjt}(K_j)$ is the ideal log-change weight for good k and $CPI(C_{kj})$ is the "conventional" price index constructed assuming the basket of varieties to be constant.
As long as the set of $\{\lambda_{kjt}\}$ is not constant, measuring prices with a conventional price index formula thus induces a bias that may be embarrassing in empirical works. However, estimating this aggregate exact price index allows to measure the price impact of changes in the variety supply or taste parameters towards varieties of imported goods.[15] In the following, we measure the gap between conventional and exact prices for 28 importing countries j using the following formula:

$$Bias_{jt} \equiv \frac{CPI(C_{kj}) - EPI(C_{kjt})}{EPI(C_{kjt})} = \left[\Pi_{k \in K_j} \left(\frac{\lambda_{kjt}}{\lambda_{kjt-1}} \right)^{\frac{w_{kjt}(K_j)}{\sigma_k - 1}} \right]^{-1} - 1 \tag{7}$$

Comparing the results across j then allows to infer the implications of the measured price biases for *relative* price levels. The previous analysis suggests the bias should be all the stronger as the number of imported varieties increases. This should notably be the case in opening up countries. As evident from Equation (6), the magnitude of the bias does not however only depend on the number of varieties but also on the value of each imported variety and its time variance, the substituability between varieties of each imported good and the weight of each good in the total consumption. Studying the extensive price effect thus requires a rigourous strategy that allows to catch all these transmission channels.[16] The empirical strategy we use to measure these biases is described in the next section.

[15] Moreover, Feenstra (1994) shows that this index is robust to the possibility that there may be more than one variety produced in each exporting country. Indeed, changes in the number of varieties produced in a country have the same effect as changes in taste parameters $\{d_{kijt}\}$.

[16] On the contrary, Broda & Weinstein (2004) only measures changes in the number of varieties in each importing country and their simulated measurement biases are thus inaccurate.

3 Empirical strategy

In the following, the bias of new varieties is estimated by applying Equation (7) to an exhaustive trade database. To this aim, we successively consider 28 importing countries (j) and study all their bilateral trade relations with the remaining 129 trade partners in each sector k in which they import a positive amount. Estimating this bias only requires to measure:

- the good- and time-specific fractions of expenditures spent in varieties imported in two consecutive periods by country j ($\{\lambda_{kjt}\}$ and $\{\lambda_{kjt-1}\}$). These shares are directly computable using observed time series of bilateral trade flows,
- the good- and time-specific "ideal log-change weights" (w_{kjt}), obtained from expenditure shares :

$$w_{kjt} \equiv \frac{\frac{s_{kjt}-s_{kjt-1}}{\ln s_{kjt}-\ln s_{kjt-1}}}{\sum_{k\in K_j}\left(\frac{s_{kjt}-s_{kjt-1}}{\ln s_{kjt}-\ln s_{kjt-1}}\right)}, \quad s_{kjt} = \frac{\sum_i p_{kijt}x_{kijt}}{\sum_k \sum_i p_{kijt}x_{kijt}}$$

- and the good-specific elasticities of substitution across varieties (σ_k).

Among these variables, only the elasticities of substitution have to be estimated. To this aim, Feenstra (1994) proposes a method that uses the reduced form of a simple model of import demand and supply. From these estimates, one can compute the importer-specific bias of new varieties and compare it across countries and over time.

3.1 A model of import demand and supply

The first step before estimating the elasticities is to precise our empirical definition of a "variety". From a theoretical point-of-view, one should rely on firm-level data to account for this concept. Indeed, in the Dixit-Stiglitz framework described in the previous section, each producer finds it profitable to produce a differentiated variety rather than mimiking an existing one. In this case, varieties are thus differentiated at the firm-level. As it is however difficult to obtain bilateral firm-level data for more than a handful of countries, the following analysis calls variety a good produced in a particular country.[17] Each industry is then assimilated to a "good" (k) whereas each bilateral trade flow in this industry identifies a particular "variety" (i). Of course, this approximation requires to use data that are as disaggregated as possible.[18] Under this definition of a variety, Feenstra (1994) proposes a method to estimate the elasticity of substitution between varieties of a good. This method uses a simple model of import demand and supply equations and applies Leamer (1981)'s result that "*if the demand and supply equations have an independent error structure, the set of possible maximum likelihood estimates of the demand and supply elasticities lie on a hyperbola defined in the second moments of the data*". To determine the point of the hyperbola defining this unique maximum

[17] Such an hypothesis is consistent with the Armington assumption.

[18] Feenstra (1994) and Broda & Weinstein (2005) work at the 8 to 10-digit level of the TSUSA nomenclature.

likelihood estimate, Feenstra uses the cross-country dimension of a panel dataset.[19] This method presents the advantage of being robust to the simultaneity bias and the measurement bias linked to the use of unit values as a proxy for prices.
Using the same notations as previously, the demand and supply system for a specific variety i at time t can be written as follows:[20]

$$\left\{ \begin{array}{lcl} p_{it} & = & exp(\nu_{it})x_{it}^{\omega} \\ x_{it} & = & \left(\frac{p_{it}}{\phi_t}\right)^{1-\sigma} \frac{d_{it}E_t}{p_{it}} \end{array} \right.$$

where

- x_{it} is the quantity of variety i bought in period t,
- $\phi_t = \left(\sum_{i \in C_t} d_{it} p_{it}^{1-\sigma}\right)^{\frac{1}{1-\sigma}}$ is the cost-minimizing function for the imported good in period t (with C_t the set of varieties available in period t and σ the elasticity of substitution between varieties),
- $E_t = \sum_{i \in C_t} p_{it} x_{it}$ is the nominal consumption,
- ν_{it} is a random technology factor assumed to be independent from d_{it},
- $\omega \geq 0$ is the inverse supply elasticity, assumed to be the same across variety suppliers.

The first equation describes the way the representative exporter from i sets its price for sales in the considered country. It is assumed to depend on the demand addressed to it and a technology parameter. The second equation explains the import demand of variety i under CES preference.

This system is then transformed to make the share s_{it} of variety i in the nominal consumption appear instead of the imported quantity. Moreover, equations are log-linearized and first-differenced. Assuming taste and technology parameters (d_{it} and ν_{it}) to be random, the previous system becomes:[21]

$$\left\{ \begin{array}{lcl} \Delta \ln s_{it} & = & \varphi_t - (\sigma - 1)\Delta \ln p_{it} + \epsilon_{it} \\ \Delta \ln p_{it} & = & \psi_t + \frac{\omega}{1+\omega\sigma}\epsilon_{it} + \delta_{it} \end{array} \right.$$

where φ_t and ψ_t are random variables that are the same whatever the variety whereas ϵ_{it} and δ_{it} are exporter-specific white noises. These variables are defined as:

$$\begin{array}{lcl} \varphi_t & \equiv & (\sigma - 1)\Delta \ln \phi_t \\ \psi_t & \equiv & \frac{\omega}{1+\omega\sigma}[\varphi_t + \Delta \ln E_t] \\ \epsilon_{it} & \equiv & \Delta \ln d_{it} \\ \delta_{it} & \equiv & \frac{1}{1+\omega\sigma}\Delta\nu_{it} \end{array}$$

[19] This method lies on the assumption that the demand and supply elasticities are good-specific and not exporter-specific.

[20] In the following, indices j and k have been suppressed for clarity. However, as the system is estimated at the industry- and country-level, all the series and estimated coefficients have a (j, k) dimension.

[21] See details in Appendix A.1.

This system serves us to estimate the elasticities of demand and supply by exploiting the panel nature of data, as explained in the following subsection.

3.2 Estimated equation

To estimate the elasticities σ and ω, the previous system is first rewritten in differences relative to a reference country r:[22]

$$\begin{cases} \Delta \ln s_{it} - \Delta \ln s_{rt} & = & -(\sigma - 1)(\Delta \ln p_{it} - \Delta \ln p_{rt}) + \tilde{\epsilon}_{it} \\ \Delta \ln p_{it} - \Delta \ln p_{rt} & = & \frac{\omega}{1+\omega}(\Delta \ln s_{it} - \Delta \ln s_{rt}) + \tilde{\delta}_{it} \end{cases}$$

then transformed into a single equation:

$$Y_{it} = \frac{\omega}{(1+\omega)(\sigma-1)} X_{1it} + \frac{(\sigma-1)\omega - (1+\omega)}{(1+\omega)(\sigma-1)} X_{2it} + u_{it} \tag{8}$$

with

- $Y_{it} = (\Delta \ln p_{it} - \Delta \ln p_{rt})^2$
- $X_{1it} = (\Delta \ln s_{it} - \Delta \ln s_{rt})^2$
- $X_{2it} = (\Delta \ln s_{it} - \Delta \ln s_{rt})(\Delta \ln p_{it} - \Delta \ln p_{rt})$
- and $u_{it} = \frac{1+\omega\sigma}{(\sigma-1)(1+\omega)} \tilde{\epsilon}_{it} \tilde{\delta}_{it}$

As shown by Feenstra (1994), ignoring the variations within each country over time, the following asymptotic relations are met:

$$E(\bar{X}_{1i}, \bar{u}_i) = 0 \ , \ E(\bar{X}_{2i}, \bar{u}_i) = 0$$

(with upper bars denoting sample means and $E(\bar{u}_i) = 0$).
$\theta_1 \equiv \frac{\omega}{(1+\omega)(\sigma-1)}$ and $\theta_2 \equiv \frac{(\sigma-1)\omega-(1+\omega)}{(1+\omega)(\sigma-1)}$ can thus be consistently estimated using the between estimator of Equation (8):

$$\bar{Y}_i = \theta_1 \bar{X}_{1i} + \theta_2 \bar{X}_{2i} + \bar{u}_i \tag{9}$$

This however requires that X_{1it} and X_{2it} are not collinear asymptotically. Feenstra (1994) shows that this is the case if the relative variance of demand and supply equations across countries are not identical :

$$\frac{\sigma^2_{\epsilon i} + \sigma^2_{\epsilon r}}{\sigma^2_{\epsilon i'} + \sigma^2_{\epsilon r}} \neq \frac{\sigma^2_{\delta i} + \sigma^2_{\delta r}}{\sigma^2_{\delta i'} + \sigma^2_{\delta r}}$$

Moreover, if equation (9) is augmented with a constant, the estimation is consistent even in the presence of measurement errors due to the use of unit values to proxy prices.[23]

[22]This country has to be a supplier of the good in each period. This constraint can be demanding in some cases, when no exporter is continuously present on the market between 1994 and 2003.

[23]As argued by Feenstra (1991), the use of unit-values "*gives motivation for treating the taste parameters for each product and supplying country as a random.*"

In the following, this model is estimated using an OLS method that uses dummy variables as instruments for the equation in levels.[24]

3.3 Computation of substitution elasticities

Estimating equation (9) provides us with consistent estimates of θ_1 and θ_2. The j- and k-specific elasticity σ is then obtained using the following formula:

- if $\hat{\theta}_1 > 0$ and $\hat{\theta}_2 > 0$ then:

$$\hat{\sigma} = 1 + \frac{1}{\hat{\theta}_2}\left(\frac{2\hat{\rho}-1}{1-\hat{\rho}}\right) \; where \; \hat{\rho} = \frac{1}{2} + \left(\frac{1}{4} - \frac{1}{4+\frac{\hat{\theta}_2}{\hat{\theta}_1}}\right)^{\frac{1}{2}}$$

- if $\hat{\theta}_1 > 0$ and $\hat{\theta}_2 < 0$ then:

$$\hat{\sigma} = 1 + \frac{1}{\hat{\theta}_2}\left(\frac{2\hat{\rho}-1}{1-\hat{\rho}}\right) \; where \; \hat{\rho} = \frac{1}{2} - \left(\frac{1}{4} - \frac{1}{4+\frac{\hat{\theta}_2}{\hat{\theta}_1}}\right)^{\frac{1}{2}}$$

If $\hat{\theta}_1 < 0$ however, one cannot obtain estimates of σ that are consistent with the theory (*i.e.* that are higher than unity). To deal with this problem, Broda & Weinstein (2005) propose to conduct a grid search to find the minimum sum of weighted least shares of the residuals of Equation (9) over a possible range of values for $\sigma (> 1)$ and ω ($\in [0; 1]$).
In the following, this method is applied to the BACI database described hereafter.

4 The Data

4.1 The Variables

To measure the price effect of new varieties, data from BACI[25] are used. This dataset provides harmonized bilateral trade flows for more than 5,000 hs6 products and 130 countries, over the 1988-2003 period. In the following, we however focus on the 1994-2003 period because of the high number of missing values before 1994. Moreover, the results are only presented for 28 countries.[26]

[24]The method is identical as the one used by Feenstra (1994). Alternatively, Broda & Weinstein (2005) estimate the model by weighted least squares, each i-specific observation being weighted by the number of adjacent time periods for which variety i is available. Both methods are strictly equivalent.

[25]*Base Analytique du Commerce International*, a bilateral trade database developed at CEPII drawing on the United Nations ComTrade database. All details concerning BACI are provided on the CEPII's website, `www.cepii.fr/francgraph/bdd/baci.htm`.

[26]Argentina, Australia, Belgium, Brazil, Canada, China, Denmark, France, Germany, Greece, Hong Kong, India, Indonesia, Italy, Japan, Korea, Mexico, Netherlands, Poland, Russian Federation, Singapore, Spain, Sweden, Switzerland, Taiwan, Turkey, the United Kingdom and the United States.

For each sector hs6 (k) in each importing country (j), estimates of θ_1 and θ_2 are obtained from equation (9), from which we get an estimation of σ. To this aim, we need to measure bilateral prices and the share of each exporter in imports of good k by country j. In the following, prices are proxied by unit values (computed at the sectorial level for each importing country):

$$p_{it} = \frac{Val_{it}}{Qty_{it}}$$

with Val_{it} the value of imports from country i and Qty_{it} the imported volume (in tons). Expenditure shares are defined as:

$$s_{it} = \frac{Val_{it}}{\sum_{i \in C_t} Val_{it}}$$

From these series, the variables Y_{it}, X_{1it} and X_{2it} entering equation (9) can be obtained and the coefficients θ_1 and θ_2 estimated. When the elasticity of substitution cannot be estimated directly, we use Broda & Weinstein (2005)'s grid search technique over values for σ between 1 and 80 and for 33 values of ω between 0 and 1. This simulation method allows us to fill in the distribution of sectorial elasticities of substitution. However, there are still cases in which σ cannot be estimated, namely when no exporter is present on the considered market during the whole period in which case there is no eligible "reference" country. In that case, the corresponding sector is ignored from the computation of the aggregate bias $Bias_{jt}$.

Once the elasticities of substitution have been estimated, one calculates the measurement bias in price indices using the analytical results of Section 2 with $t-1 = 1994$ and $t = 2003$. The aggregate bias is thus calculated for the whole 1994-2003 period. However, it will sometimes be convenient to annualize results using the following formula: $Bias^y_{jt} \equiv (Bias_{jt}+1)^{1/n}-1$ with n the length of the period (9 years in most cases).

To compute $Bias_{jt}$, the shares λ_{kjt} (λ_{kjt-1}) of the sectorial imports in 2003 (1994) carrying on varieties imported in both years are first calculated. The ratio $\lambda_{kjt}/\lambda_{kjt-1}$ raised at the power $1/(\hat{\sigma}-1)$ measures the sectorial bias induced when ignoring new varieties of the good k to build price index. Finally, the whole distribution of sectorial biases is aggregated using Equation (7) to obtain a measure of the bias that can be interpreted from a macroeconomic viewpoint. This aggregation involves the matrix of weights $\{w_{kjt}(K_j)\}$:

$$w_{kjt} = \frac{\frac{s_{kjt}-s_{kjt-1}}{\ln s_{kjt}-\ln s_{kjt-1}}}{\sum_{k \in K_j} \frac{s_{kjt}-s_{kjt-1}}{\ln s_{kjt}-\ln s_{kjt-1}}}$$

with

$$s_{kjt} = \frac{\sum_{i \in C_{kjt}} Val_{kijt}}{\sum_{k \in K_j} \sum_{i \in C_{kjt}} Val_{kijt}}$$

Consistent with the theoretical model, we assume that the basket of goods imported by country j (K_j) is the same in $t-1$ and t. However, this is not necessarily true. In the case we cannot identify a set C_{kj} of varieties that are imported in 1994 and in 2003, we switch to the higher aggregation level, that is the hs4 level. In this case, all bilateral flows in a given hs4 category are considered as imperfect substitutes of the same good k.

Once estimated the distribution of elasticities, measuring the bias of endogenous variety supply is thus easy. Before moving to the results, it is however interesting to get an idea of the extensive effect phenomena in BACI.

4.2 Magnitude of extensive margins in BACI

To get an idea of the size of extensive margins in the BACI database, the annual growth of the value of good k imported by country j is decomposed into three parts:[27]

- the intensive growth, measured by the percentage change in the value of varieties continuously imported between $t-1$ and t,
- the value of "new" varieties that country j was not importing in year $t-1$, normalized by the total value of imports in $t-1$,
- and the imported value of varieties that disappeared between $t-1$ and t, also divided by imports in $t-1$.

The difference between the second and third components is what we call the extensive margin. This decomposition is done for each sector k and each importer j for the results not to be biased by changes in the set of importing countries or the nomenclature of products occuring during the period under consideration.

A summary of the results is illustrated in Figure 1 that sums up each component on the whole set of products k and countries j.[28] This gives us a decomposition of the annual growth of the total value of imports.[29] On average on the whole period, the extensive margin (difference between the growth rates in dark and light grey in Figure 1) only explains 0.15 percentage point of the annual growth of world imports. Indeed, the value of appearing flows more or less compensate for the value of disappearing ones. However, decomposing the growth of the imported value tends to minimize the size of the extensive margin. When studying the *number* of bilateral flows, the magnitude of this phenomenon get higher: while "extensive" flows count for only 3.9% of the value of imports between 1994 and 2003, they represent 40.1% of observations. This simply reflects the fact that new flows are generally of small size in their first year of apparition. Moreover, using $hs6$ data and assimilating each exporting country as a differenciated variety mechanically leads to underestimate the extensive margins of trade as goods produced in a given sector by a specific exporting country are themselves differentiated.

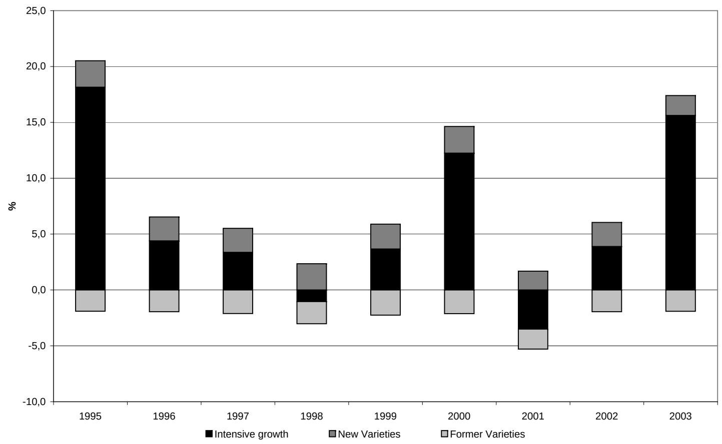
Whereas the aggregate nominal effect is weak, the extensive margin can still be an important share of import growth at the sectorial level. Table A.1 in Appendix A.2.2 gives an equivalent decomposition of the 1994-2003 growth of world imports at the 2-digit industrial level.

[27] Details on this decomposition are provided in Appendix A.2.1.

[28] Figure 1 is constructed without selecting countries as done in the rest of the paper.

[29] The total growth rate does not correspond exactly to the true annual growth rate of world trade during the period under consideration because some filters have been applied to BACI for results not to be biased by problems linked to the quality of data. These treatments are described in Appendix A.2.1.

Figure 1: Decomposition of the annual growth of bilateral flows (in %)



Source: BACI.

This table shows that the magnitude of the extensive margin strongly varies across sectors.[30] Indeed, more than half the growth of imports is due to extensive effects in the following categories: 54 ("Man-made filaments"), 89 ("Ships, boats and floating structures") and 97 ("Works of art, collectors' pieces and antiques").

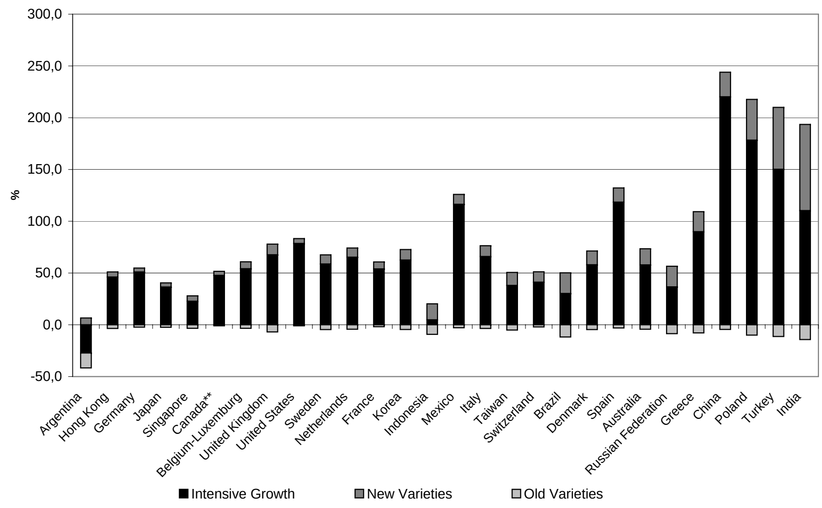
Discrepancies across importing countries are also significant as shown in Figure 2.[31] In this figure, countries are ranked depending on the size of the extensive margin during the whole period. Among the considered importing countries, Argentina if the one with the weakest extensive margin. Indeed, the value of disappearing flows between 1994 and 2003 is higher than the value of appearing flows and the extensive margin has a negative effect on import growth. On the contrary, the phenomenon is especially strong in India and Indonesia where the extensive components explain more than 35% of the import growth (more than 50% in Indonesia). This graphic shows a positive correlation between the magnitude of the extensive effect and the total growth of imports.[32] Such a correlation is consistent with the explanation of the phenomenon in Mélitz (2003) type models: the market potential of fast growing countries (in terms of activity then of imports) increases quickly, which strengthens the incentive for foreign producers to pay the fixed cost for entering this market through

[30] Of course, this heterogeneity would be even stronger at a more disaggregated level.

[31] In this figure, the total import growth between 1994 and 2003 is studied, except for Belgium (1995-2003), Canada (1994-2002) and the Russian Federation (1996-2003). These restrictions are justified in Appendix A.2.1.

[32] The correlation coefficient among the sample of 28 importing countries is 0.738.

Figure 2: Decomposition of the 1994-2003 import growth, by importer



Source: BACI

export.[33]

In some given sectors or countries, the extensive effects thus contribute to a large share of the import growth. As a consequence, they are likely to have a significant effect on prices. Moreover, the cross-country heterogeneity suggests that the extensive price effect should vary depending on the considered country. As a consequence, the error in measured price indices should transmit into relative price levels.

5 The Results

In the following, the results obtained at each stage of the empirical analysis are detailed successively. First, a quick description of the sectorial elasticities of substitution estimated at the product- and country-level is provided. Then, we compute the sectorial measurement biases, before aggregating them at the country-level. This provides us with a synthetic measure of the bias in conventional import price indices that can be compared across importing countries.

[33]Under this interpretation, the negative extensive effect obtained for Argentina can be attributed to the 1998-2002 crisis that may have lead exporters getting away from this market.

5.1 Estimated elasticities of substitution

Estimates of σ at the product- and country-level are illustrated in Figure 3 and detailled in Table A.2 in Appendix A.3. This estimation is systematically conducted at the hs6 level using the strategy described in Section 3. When necessary, these elasticities are however aggregated at the hs4 or hs2 level by weighted mean calculations. At the bottom of Table A.2 (the line labelled "BW SITC-5, 90-01"), results obtained by Broda & Weinstein (2005) on US data are provided for comparison.[34]

On average, these estimates suggest a strong substituability between varieties in each sector: on average over all countries, the median estimate is 6.2. Mean and median coefficients are indeed pushed upwards by a number of very high elasticities. Comparing these results with other estimations of this parameter is however tricky, the empirical literature being unable to reach a consensus. Some papers infer this elasticity from mark-up data, like Morrison (1990) that chooses a value of 6, close to our median estimation. Our distributions of estimated coefficients are, however, more concentrated on high σ values than those obtained by Broda & Weinstein (2005) using the same empirical strategy.[35]

These distributions are also heterogeneous in the country dimension. The median elasticity thus varies between 4.6 in India and 10.8 for Mexico. This heterogeneity can be explained by a composition effect and/or an aggregation bias. Indeed, as countries do not import the same basket of goods, these statistics are affected by composition effects. Moreover, the use of the hs6 dimension of data to define a "good" induces an aggregation bias which effect may vary among countries: products grouped in a single hs6 sector are not necessarily the same from an importing country as another one which can explain discrepancies in elasticities at the sectorial level.

With respect to Broda & Weinstein (2005), these high elasticites should lead to an underestimation of the aggregated measurement bias obtained when computing equation (6). As a consequence, they are used carefully in the rest of the paper: the sectorial biases presented in the next section do not take into account their impact and aggregate biases are successively calculated using different distributions of σ coefficients, the estimated ones as well as several uniform distributions.

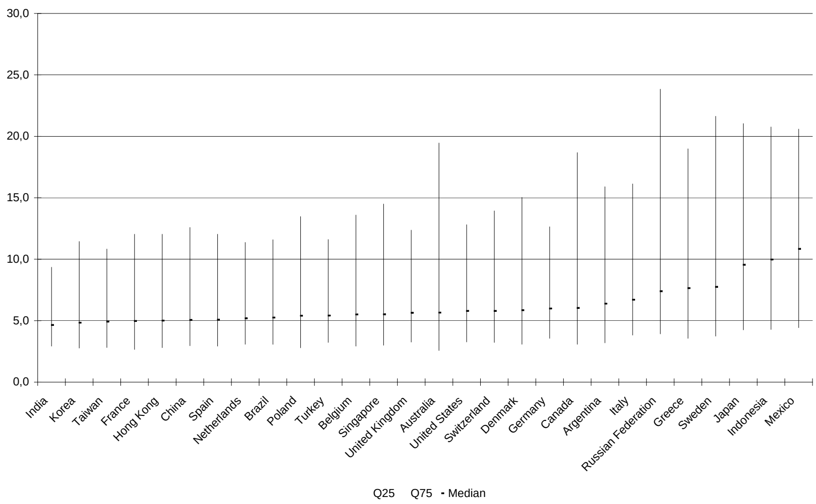
5.2 Sectorial bias in measured prices

Figure 4 completed with Table A.3 in Appendix A.3 provides descriptive statistics on the measurement biases computed at the sectorial level. Here again, Broda & Weinstein's results are given in the last line of Table A.3 for comparison reasons.

[34]Results are those obtained by Broda & Weinstein (2005) at the SITC-5 level because this classification is closer to the hs6 one used in this paper than the $TS10$ classification which is the finest in use in Broda and Weinstein's paper.

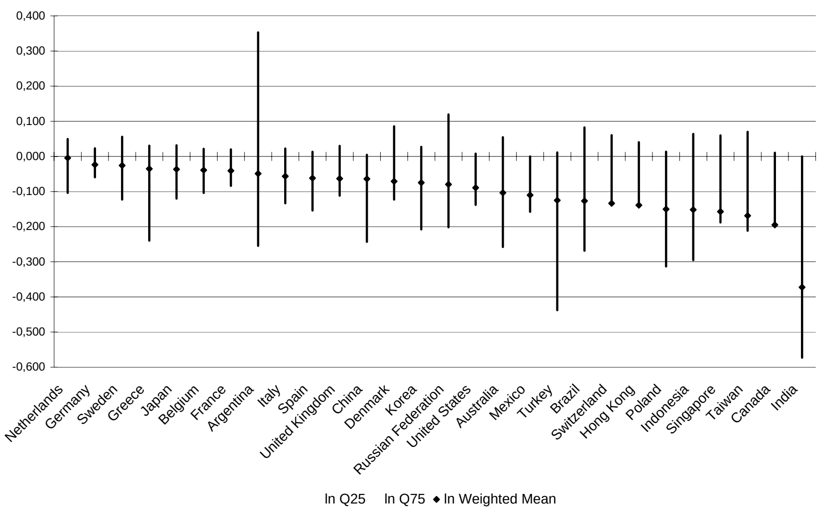
[35]These gaps are partly linked to the different datasets used in both papers. Indeed, when using our method on the NBER database described by Feenstra (1996), one obtains elasticities between those of Broda & Weinstein (2005) and ours. On average: the median estimate is 3.2 and the mean coefficient is 8.6.

Figure 3: Descriptive statistics (median, first and third quartiles) of the estimated distributions of substitution elasticities (period 1994-2003)



Source: Estimations from BACI

Figure 4: Descriptive statistics (weighted geometric mean, first and third quartiles, in log) concerning the distributions of ratios $\lambda_{kjt}/\lambda_{kjt-1}$ (period 1994-2003)



Source: BACI

As expected, the ratios $\lambda_{kjt}/\lambda_{kjt-1}$ are on average lower than unity, which means that the share, in total imports, of varieties consumed in 1994 and in 2003 tends to decrease over time. This reflects an expenditure switching towards new varieties operating during the considered period. On average, measuring sectorial prices without taking into account changes in the variety supply leads to an overestimation of the true price index of around 9%.[36]

As expected, the distribution of sectorial biases strongly varies among importing countries. It is more concentrated on weak values (that imply a strongly positive measurement bias) in emerging countries and Canada. These results are consistent with our analysis of extensive effects in section 4.2: the correlation between the mean $\lambda_{kjt}/\lambda_{kjt-1}$ ratio and the magnitude of the extensive growth illustrated in Figure 2 is -0.63. India is the country in which biases are the higher: on average, a sectorial price index that neglects new varieties overestimate price growth by more than 30 % between 1994 and 2003.

Comparing results for the US with those obtained by Broda & Weinstein (2005) with the NBER database, one observes that their distribution of ratios is more spread out than ours,

[36]This figure is of course very approximative as it does not take into account the role of substitution elasticities but corresponds to the weighted mean of the observed $\lambda_{kjt}/\lambda_{kjt-1}$ ratios. It is thus implicitely assumed that σ is constant and equal to 2, in which case the sectorial bias is: $(\lambda_{kjt}/\lambda_{kjt-1})^{1/(\sigma-1)} = \lambda_{kjt}/\lambda_{kjt-1}$.

which implies stronger (positive or negative) biases: whereas the range between the 5th and the 95th percentile is 2.55 in Broda & Weinstein (2005), it is only 0.90 in our estimates.[37] Several explanations of this discrepancy are conceivable. First, it must be noted that the true discrepancy is smaller than suggested in Table A.3. Indeed, Broda & Weinstein's results cover the 1990-2001 period, *i.e.* 11 years whereas our estimates span on 9 years. When rescaled to a comparable period, the range between the 5th and the 95th percentile diminishes to 1.99 in Broda & Weinstein's estimates. Still however, our biases are weak on average. One conceivable explanation lies in the treatments of trade data in BACI. Indeed, the reconciliation of import and export declarations complete the bilateral trade matrix, thus reducing the volatility in the number of bilateral flows. This treatment can of course impact on the measure of extensive effects. As an illustration, imagine that the United States ceases declaring imports from a given country during the considered period. In the NBER database, this event will lead to the disappearance of several varieties whereas this will not necessarily be the case in BACI, if the exporting country still declares the amount it sells in the United States. The bias attributable to the quality of data should thus be weaker in estimates obtained from BACI. However, data provided by the United States are *a priori* of good quality and this quality bias should not be too pronounced. Another conceivable explanation of the relative weakness of our biases concerns the period of estimation which is more recent than Broda & Weinstein's. Now, these authors' results suggest that in the United States the price measurement bias is higher on the 1972-1988 period than in the 1990-2001 period. If this decreasing trend has gone on, the measurement bias should indeed be weaker on the 1994-2003 period.

Even if the US bias is on average weaker than in Broda & Weinstein, these biases are still important in some sectors or countries. Moreover, the possible underestimation is less of a problem when conducting a cross-country comparison, at least if this problem is spread across all countries. In the following, these sectorial biases are thus aggregated using equation (6) so as to obtain a macroeconomic measure of the bias of omitted new varieties.

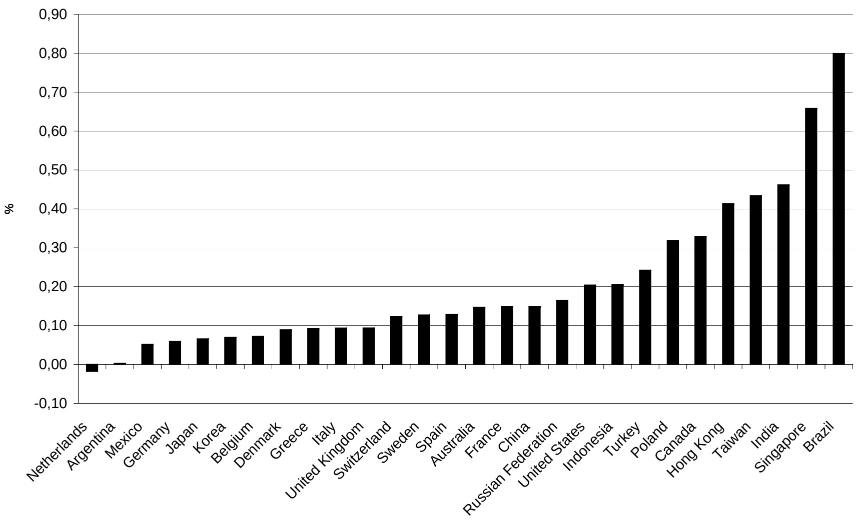
5.3 Aggregated measurement bias

Figure 5 and Table A.4 in Appendix A.3 allow to compare across importing countries the aggregate price measurement bias due to the omission of endogeneous changes in the supply of variety. These aggregate biases are constructed from the previously described $\lambda_{kjt}/\lambda_{kjt-1}$ ratios and using equations (6) and (7).[38] As expected, this bias is generally positive: neglecting new varieties leads to overestimate true import prices. This bias is weak on average, around .2% per year, but exceeds .4% a year in several Asian emerging countries and even .8% in Brazil.

[37] This comparison uses Broda & Weinstein's results at the $SITC5$ sectorial level, which is the most comparable with our $hs6$ nomenclature. The range between the 5th and the 95th percentile of their distribution of $TS10$ sectorial biases is smaller (2.24) but still larger than ours.

[38] We have however dropped those $\lambda_{kjt}/\lambda_{kjt-1}$ ratios describing variety supply changes in the hs2 category called "Mineral fuels, oils & product of their distillation". Indeed, trade in these sectors is highly specific and as it sometimes accounts for a large share of a country's imports, it could be misleading to incorporate it in the computed measurement biases.

Figure 5: Annual measurement bias of prices, in %



Source: BACI

These cross-country discrepancies are not perfectly correlated with gaps in the extensive margin intensity (Figure 2) or the mean sectorial bias (Figure 4). Indeed, the aggregate measure of the price bias is influenced by the weight of each sector in the import basket as well as the distribution of estimated elasticities described in Section 5.1. Thus, a low $\lambda_{kjt}/\lambda_{kjt-1}$ ratio has a higher impact on the aggregate bias if country j imports a lot of good k as well as if the elasticity of substitution estimated from the (j, k) sub-sample is weak. To have an idea of the role of the estimated σs in explaining cross-country discrepancies in terms of aggregate price bias, the last four columns of Table A.4 report the bias obtained when constraining the elasticity of substitution to be constant across sectors.[39] One observes that the relative weakness of our estimated biases is partly attributable to the strength of the estimated elasticities. When σ is calibrated to reproduce Broda & Weinstein's mean results for the United States (column "$\sigma = 2.66$" of Table A.4), the measurement bias increases in almost all countries[40] and reaches 1.4% a year in India.

Comparing the results across importers shows that the price measurement bias is the highest in fast growing countries, notably in Asia, whereas it is lower in rich countries. Thus, the correlation between these biases and the growth of GDP[41] is 50% (with a p-value of 0.006).

[39] Successively $\sigma = 2$, $\sigma = 2.66$, $\sigma = 3$ and $\sigma = 5$ the second value being calibrated to reproduce for the United States the bias obtained by Broda & Weinstein (2005).

[40] It decreases, however, in Brazil, China and Netherlands.

[41] Source: *Penn World Tables*, Heston, Summers & Aten (2002). Mean annual growth rate computed

Mélitz (2003) type models of heterogeneous firms provide a theoretical framework to understand this correlation. In these models, firms entry decisions in export markets are determined by the expected profits in each national market, that notably depend on the country's market potential. The fast growth of emerging countries that increases their demand of imported goods could thus have conducted new exporters to enter these markets, with a decreasing impact on import prices.
Several countries are, however, outliers. The Canadian bias is relatively high because of a composition effect: the weighted mean of sectorial biases is pushed downwards by a few large sectors.[42] On the contrary, the aggregate bias is surprisingly weak in Korea, Mexico and, above all, Argentina. With respect to Korea and Mexico, this is for a large part due to the weakness of estimated elasticities in sectors where the $\lambda_{kjt}/\lambda_{kjt-1}$ ratio is weak. Indeed, when constraining the elasticity of substitution to be uniform across sectors, the position of these countries in the distribution of aggregate biases is rather central (see Figure A.1 in Appendix). Moreover, in the Mexican case, the use of sectoral rather than firm-level data may be misleading as this country's imports are strongly dominated by US products. Thus, the growth of import variety is probably mainly reflected in the increasing number of US firms entering this national market, which are not marked as new varieties in our empirical strategy. Last, the weakness of the Argentinian bias remains even when constraining the secdtorial σs to be the same. Indeed, this result can be explained by the number of negative extensive effects leading to $\lambda_{kjt}/\lambda_{kjt-1}$ ratios that are greater than unity. This is probably a consequence of the crisis that strucked this country in the beginning of the 21st century, which probably pushed numerous exporters out of the market.

These results thus suggest that standard price series that do not take into account changes in the variety supply of imported goods tend to overestimate the true price index, notably in emerging countries. This cross-country heterogeneity has important consequences in terms of relative prices. As an illustration, consider Table A.5 in Appendix that gives the estimated bias in the relative import price index with respect to the United States, attributable to the omission of changes in the supply of variety available in each market. As expected from Figure 5, this bias in relative prices is negative in most countries, but positive for several emerging countries[43] and Canada, meaning that relative import prices with respect to the US are overestimated in those countries. As shown by the comparison in Table A.5 of column 2, that gives the estimated bias in relative import prices, and column 3, that provides us with the growth of the "conventional" relative import price series computed by the IMF,[44] this bias in relative prices may be sizeable in some countries. For instance, neglecting the price impact of changes in the relative supply of varieties in Switzerland and the United States

from the real GDP in PPP on the 1994-2000 period.

[42] A more detailed analysis of the Canadian results proves that this composition effect is mainly due to four hs2 catehories: 44. "Wood and Articles of wood" which counts for 1% of the total imports and which $\lambda_{kjt}/\lambda_{kjt-1}$ ratio is 0.8, 49. "Printed books, newspapers, pictures" (1.1% of imports and a ratio of 0.53), 88. "Aircraft, Spacecraft" (1.3% of imports and a ratio of 0.66) and 90. "Optical, photo, cine, meas, checking, precision, etc" (1.9% of imports and a ratio of 0.84).

[43] Namely, Indonesia, Turkey, Poland, Hong Kong, Taiwan, India, Singapore and Brazil.

[44] Source: *International Financial Statistics*, Import unit value index.

leads to underestimate the growth of the Swiss relative import price by 70%. On the contrary, whereas the IMF estimates that the Polish import price has increased by 4.8% with respect to the US, taking into account firms' entry decisions would reduce this figure to 3.8%. This suggests that neglecting changes in the variety available in national markets is able to generate an import measurement error in relative prices.

6 Conclusion

By using a method first proposed by Feenstra (1994) and applying it to the BACI database, this paper compares the price impact of endogenous changes in the variety supply of imported goods, over a sample of 28 countries. Following Broda & Weinstein (2005), the empirical strategy lies on the idea that "conventional" price indices are mismeasured because they do not integer immediately newly imported varieties of a good. According to the New Trade literature, these "extensive" fluctuations have yet an impact on aggregate prices when consumers value variety. Based on this intuition, it is thus interesting to measure the aggregate price impact of newly imported varieties and compare it across countries. To this aim, the BACI database is especially well adapted because it covers bilateral trade at a highly disaggregated level with an exhaustive country coverage. Assuming that each exporting country of a given good produces a differentiated variety, one can thus apply Feenstra's method to a large range of goods and countries.

Our estimations confirm the downward influence of endogenous changes in variety supplies on import price levels. On average between 1994 and 2003, the appearance of new varieties in the considered import markets lead to an unrecorded 0.2% annual drop in import prices. Moreover, our analysis shows that this price impact strongly varies across sectors and countries. The measurement bias is thus especially pronounced in emerging countries, notably in Asia. Our interpretation of this finding is that fast growing countries tend to attract new exporters with their high market potential ; this extends these countries' variety supply of imported goods and reduces their price level.

These results have important consequences in terms of relative prices. Indeed, they suggest that firms' entry in international markets have a real depreciating effect in countries benefiting from an expansion of the variety of goods they can consume. As "conventional" price indices built on a constant basket of products neglect this effect, measured relative prices could thus be overestimated, notably in fast growing countries that tend to attract exporting firms. This mismeasured effect is all the more of a problem since it is likely to counteract the Balassa-Samuelson effect. Indeed, those productivity gains that lead to a real appreciation in the standard Harrod-Balassa-Samuelson model[45] are also likely to enhance the country's attractiveness for foreign exporters. The following entry of firms in this country's import markets should impact its relative price in the opposite direction as the Balassa-Samuelson effect.

The magnitude of the measured effects is however weak on average. However, our results can be considered as lower bounds for at least two reasons. First, the use of sectorial rather than

[45] See Harrod (1933), Balassa (1964) and Samuelson (1964).

firm-level data leads to underestimate extensive effects because individual firms entering a market in which some of their national competitors are already present are not recorded in the extensive margin. Secondly, using trade data that neglect the domestic supply of differentiated varieties could also introduce a bias in the measured price effects. Indeed, according to Helpman, Melitz & Yeaple (2004), firms decisions to enter a market through export are only an intermediate stage before doing FDI. As a consequence, a sustained growth of the demand emanating from a (developing) country should encourage firms to invest in production units that serve the local market directly rather than exporting. Entering FDI flows and their impact on the local variety supply should thus also push relative prices downwards.

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A.1 Derivation of the estimated equation

To derive the estimated equation allowing to obtain a measure of the elasticity of substitution at the sectorial level, we start with the following import demand/supply system:

$$\left\{\begin{array}{lcl} p_{it} & = & exp(\nu_{it})x_{it}^{\omega} \\ x_{it} & = & \left(\frac{p_{it}}{\phi_t}\right)^{1-\sigma}\frac{d_{it}E_t}{p_{it}} \end{array}\right.$$

This system is then log-linearized and transformed to make the share s_{it} of variety i in the nominal consumption appear instead of the imported quantity:

$$\left\{\begin{array}{lcl} \ln p_{it} & = & \frac{1}{1+\omega}\left[\nu_{it} + \omega \ln s_{it} + \omega \ln E_t\right] \\ \ln s_{it} & = & (1-\sigma)\ln p_{it} + (\sigma-1)\ln\phi_t + \ln d_{it} \end{array}\right.$$

Assuming taste and technology parameters (d_{it} and ν_{it}) to be random, the previous system becomes in first-differences:

$$\left\{\begin{array}{lcl} \Delta\ln s_{it} & = & \varphi_t - (\sigma-1)\Delta\ln p_{it} + \epsilon_{it} \\ \Delta\ln p_{it} & = & \psi_t + \frac{\omega}{1+\omega\sigma}\epsilon_{it} + \delta_{it} \end{array}\right.$$

where φ_t and ψ_t are random variables that are the same whatever the variety whereas ϵ_{it} and δ_{it} are exporter-specific white noises. These variables are defined as:

$$\begin{array}{lcl} \varphi_t & \equiv & (\sigma-1)\Delta\ln\phi_t \\ \psi_t & \equiv & \frac{\omega}{1+\omega\sigma}[\varphi_t + \Delta\ln E_t] \\ \epsilon_{it} & \equiv & \Delta\ln d_{it} \\ \delta_{it} & \equiv & \frac{1}{1+\omega\sigma}\Delta\nu_{it} \end{array}$$

In differences relative to a reference country r, we get:

$$\left\{\begin{array}{lcl} \Delta\ln s_{it} - \Delta\ln s_{rt} & = & -(\sigma-1)(\Delta\ln p_{it} - \Delta\ln p_{rt}) + \tilde{\epsilon}_{it} \\ \Delta\ln p_{it} - \Delta\ln p_{rt} & = & \frac{\omega}{1+\omega}(\Delta\ln s_{it} - \Delta\ln s_{rt}) + \tilde{\delta}_{it} \end{array}\right.$$

which is transformed into a single equation by multiplying both equations and rearranging:

$$\begin{array}{lcl} (\Delta\ln p_{it} - \Delta\ln p_{rt})^2 & = & \frac{\omega}{(1+\omega)(\sigma-1)}(\Delta\ln s_{it} - \Delta\ln s_{rt})^2 + \frac{(\sigma-1)\omega-(1+\omega)}{(1+\omega)(\sigma-1)} \\ & & (\Delta\ln s_{it} - \Delta\ln s_{rt})(\Delta\ln p_{it} - \Delta\ln p_{rt}) + \frac{1}{\sigma-1}\tilde{\epsilon}_{it}\tilde{\delta}_{it} \end{array}$$

A.2 Decomposition of import growth in BACI

A.2.1 Method

The import growth decomposition presented in Section 4.2 is done at the product- and country-level (j,k) using the following formula:

$$g_{jkt} = g_{jkt}^{Int} + g_{jkt}^{Ext}$$

where:

$$g_{jkt} \equiv \frac{\sum_{i \in C_{kjt}} Val_{ijkt} - \sum_{i \in C_{kjt-1}} Val_{ijkt-1}}{\sum_{i \in C_{kjt-1}} Val_{ijkt-1}}$$

is the total import growth in goods k by country j between $t-1$ and t,

$$g_{jkt}^{Int} \equiv \frac{\sum_{i \in C_{jk}} (Val_{ijkt} - Val_{ijkt-1})}{\sum_{i \in C_{jkt-1}} Val_{ijkt-1}}$$

is the "intensive" growth, on bilateral flows that are strictly positive in $t-1$ *and* t (called C_{jk} in Section 2),

$$g_{jkt}^{Ext} = \frac{\sum_{i \in new} Val_{ijkt} - \sum_{i \in old} Val_{ijkt-1}}{\sum_{i \in C_{kjt-1}} Val_{ijkt-1}}$$

is the "extensive" growth, difference between the value of newly imported varieties in t (the set $new \equiv C_{kj} - C_{kjt}$) and the value of $t-1$ imports on varieties that are no more imported in t (the set $old = C_{kjt-1} - C_{kj}$).

This decomposition is first realized on a year by year basis and used to construct Figure 1 that just sums up each component (g_{jkt}, g_{jkt}^{Int} and g_{jkt}^{Ext}) on all pairs (j, k). We also decompose the total growth of imports between 1994 and 2003, using only the information on these two years, to build Figure 2 as well as the sectorial results of Table A.1 hereafter.

For three countries, Belgium, Russian Federation and Canada, this decomposition as well as the computation of the price measurement bias lies on a different period. For these countries, studying the extensive price effect on the 1994-2003 period is indeed cumbersome. Thus, Belgium begins declaring its international trade towards the United Nations in 1995. As a consequence, if we start the analysis in 1994, the import value of new varieties will be overestimated in 1995. Estimations concerning this country thus lie on the 1995-2003 period. The same problem arises with the Russian Federation that begins declaring trade in 1996. This country's imports are thus studied on the 1996-2003 period. Last, we only consider Canadian imports between 1994 and 2002. Indeed, the number of recorded flows concerning Canada increases in an unexplained way in 2003 in the ComTrade database, probably because of a change in the customs nomenclature. Working on the 1994-2003 period would thus lead to an overestimation of the extensive margin in 2003.

In the decomposition of import growth, we tried to limit the excessive volatility in the number of flows, attributable to the quality of data. Indeed, this volatility would lead to overestimate the extensive effects. To this aim, we ran several treatments on the initial data:

- we dropt exporting and importing countries for which the annual number of bilateral flows is less than 100 as well as products traded less than 100 times during a given year,
- in period t, only those bilateral flows for which both countries have not changed their declaring status between $t-1$ and t are recorded. Indeed, given the harmonization procedure in BACI,[46] a change in the declaring status could lead to an artificial variation in the value of recorded flows, that would bias the decomposition,

[46] In BACI, when a country does not report a bilateral flow, his counterpart report is used instead.

- we only keep those pairs (importer*product) for which there is at least one observation (variety) in $t-1$ and in t, *i.e.* goods that are continuously imported by country j. This allows us to release from problems linked to changes in customs nomenclatures.

A.2.2 Sectoral decomposition of import growth

Table A.1: Decomposition of the 1994-2003 growth at the hs2 level

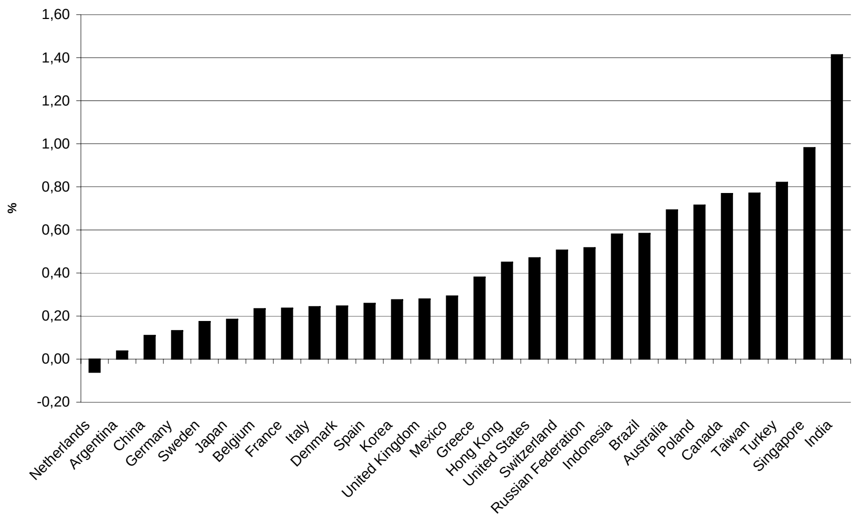
hs2 category	Extensive Growth	Intensive Growth
01. Live animals	-4.2	11.2
50. Silk	-1.8	-37.5
91. Clocks and watches	0.4	4.7
93. Arms and ammunition	0.4	2.7
41. Raw hides and skins	1.1	24.7
88. Aircraft, spacecraft	1.2	42.8
80. Tin and articles thereof.	1.2	0.4
51. Wool, fine/coarse animal hair	1.4	-16.2
67. Prepared feathers	1.9	16.2
66. Umbrellas, walking-sticks, etc.	2.1	-5.8
37. Photographic or cinematographic goods	2.3	11.1
43. Furskins and artificaial fur	2.4	26.9
06. Live tree and other plants	2.7	59.2
95. Toys, games and sports requisites	2.8	49.7
48. Paper and paperboard	3.0	55.0
53. Other vegetable textile fibres	3.2	-8.2
46. Manufactures of straw, esparto, etc.	3.3	34.2
05. Products of animal origin, nes	3.3	14.8
02. Meat and edible meat offal.	3.3	27.9
11. Prod. mill industry	3.6	17.2
42. Articles of leather	3.7	27.4
22. Beverages, spirits and vinegar	3.8	64.6
47. Pulp of wood, other fibrous cellulosic mat.	4.0	25.2
14. Vegetable plaiting materials	4.1	-4.1
07. Edible vegetables and certain roots	4.1	46.6
09. Coffee, tea, spices	4.5	-20.1
49. Printed books, newspapers, pictures	4.7	41.3
55. Man-made staple fibres	4.9	-10.2
74. Coper and articles thereof.	4.9	26.7
45. Cork and articles of cork	5.1	83.7
64. Footwear, gaiters and the like	5.1	27.9
57. Carpets and other textile floor covreings	5.2	8.9
96. Miscellaneous manufactured articles	5.3	31.0

03. Fish and crustacean	5.4	27.9
32. Tanning/dyeing extract. tannins	5.4	49.9
87. Vehicles O/t railw/tramw roll-stock, etc.	5.5	80.5
10. Cereals	5.6	7.0
24. Tobacco and manufactured tobacco substitutes	6.0	-3.3
62. Art of apparel and clothing access	6.3	36.4
82. Tool, implement, cutlery, spoon and fork	6.4	54.5
94. Furniture, bedding, mattress, etc.	6.4	96.2
84. Nuclear reactors, boilers, etc	6.5	69.1
92. Musical instruments	7.0	27.9
69. Ceramic products	7.0	28.8
52. Cotton	7.1	15.7
16. Prep of meat, fish or crustaceans	7.2	36.1
90. Optical, photo, checking, precision, etc	7.3	87.8
23. Residues and Waste from the food industry	7.3	32.0
04. Dairy prod. birds' eggs, natural honey	7.4	33.0
83. Miscellaneous articles of base metal	7.5	72.5
38. Miscellaneous chemical products	7.6	67.6
33. Essential oils and resinoids	7.6	109.2
65. Headgear and parts thereof.	7.7	55.2
85. Electrical mchy equip parts thereof.	7.7	89.5
21. Miscellaneous edible preparations	7.7	61.4
44. Wood and articles of wood	7.7	20.1
39. Plastics and articles thereof.	7.8	75.5
40. Rubber and articles thereof.	7.9	58.2
78. Lead and articles thereof.	8.2	15.9
68. Articles of stone, plaster, cement, etc.	8.9	44.8
79. Zinc and articles thereof.	9.0	29.4
35. Matières albuminoïdes	9.1	63.7
73. Ouvrages en fonte, fer ou acier	9.2	55.5
34. Savons, agents de surface organiques	9.3	83.9
08. Fruits comestibles; écorces d'agrumes ou de melons	9.4	43.5
76. Aluminium et ouvrages en aluminium	9.5	60.3
20. Préparations de légumes, de fruits	9.6	40.0
19. Préparations à base de céréales, de fécules ou de lait	9.8	82.4
17. Sucres et sucreries	10.4	17.1
28. Produits chimiques inorganiques	10.6	38.8
15. Graisses et huiles animales ou végétales	10.6	48.2
13. Lac. gums, resins and other vegetable saps	10.7	42.2
25. Salt. sulphur. earth and ston. plasering mat.	11.0	18.1
70. Glass and glassware	11.2	58.1
59. Impregnated, coated, cover/laminated textile	11.3	38.6
56. Wadding, felt and nonwoven. yarns. etc.	11.3	44.8

60. Knitted or crocheted fabrics	11.3	54.5
61. Art of apparel and clothing access, knitted or crocheted	11.4	69.5
30. Pharmaceutical products	11.4	275.1
54. Man-made filaments	12.0	3.6
58. Special woven fab. tufted textile fab. etc.	12.1	40.2
18. Cocoa and cocoa preparations	12.7	58.5
63. Other made up textile articles	12.8	74.0
26. Ores, slag and ash.	13.9	41.9
36. Explosives, pyrotechnic products, etc.	14.9	40.3
72. Iron and steel	15.6	50.2
71. Natural/cultured pearls, precious stones, etc	16.7	52.8
29. Organic chemicals	16.9	76.6
97. Works of art, collectors' pieces and antiques	17.3	16.8
31. Fertilisers	19.3	25.0
12. Oil seed, oleagi fruits, etc	20.6	37.6
86. Railw/tramw locom, rolling-stock, etc	22.3	28.2
81. Other base metals, cermets, etc.	23.5	40.0
27. Minerals fuels, oils, etc.	26.0	111.9
75. Nickel and articles thereof.	32.2	101.5
89. Ships, boats and floating structures	37.2	21.7

A.3 Additional empirical results

Figure A.1: Annual bias of price measure, in % if $\sigma_{jk} = 2.66, \forall j, k$



Source: BACI

Table A.2: Descriptive statistics concerning the elasticities of substitution estimated at the sectoral level on the 1994-2003 period

Importer	Nb of products	Elasticities of substitution: Mean	Weighted Mean$^{(a)}$	P10	Median	P90
Argentina	455	14.8	50.7	2.1	6.4	37.0
Australia	1333	23.3	35.4	1.8	5.7	50.9
Belgium$^{(b)}$	2939	15.0	15.0	1.9	5.5	35.2
Brazil	1042	14.6	18.3	2.1	5.3	31.7
Canada$^{(c)}$	2613	22.9	31.6	2.0	6.0	54.6
China	2369	15.1	16.5	2.1	5.1	35.2
Denmark	1618	17.2	21.2	2.1	5.8	37.0
France	3696	15.7	15.8	1.8	5.0	37.0
Germany	3671	17.9	17.7	2.3	6.0	31.9
Greece	1317	16.4	28.4	2.3	7.6	35.0
Hong Kong	2526	28.1	16.7	2.0	5.0	39.4
India	713	17.8	27.9	2.2	4.6	20.6
Indonesia	1250	16.1	16.0	2.7	10.0	35.2
Italy	3059	23.4	23.4	2.5	6.7	41.2
Japan	3136	19.5	21.0	2.8	9.5	33.7
Korea	1779	14.3	19.1	2.0	4.8	30.4
Mexico	2984	15.2	14.4	2.7	10.8	31.3
Netherlands	3102	18.9	16.7	2.2	5.2	31.9
Poland	1334	16.9	15.9	1.8	5.4	38.8
Russian Federation$^{(d)}$	1680	34.3	18.6	2.6	7.4	49.5
Singapoor	1645	17.3	12.6	2.1	5.5	40.8
Spain	2997	17.0	21.1	2.0	5.1	34.5
Sweden	1936	17.9	19.0	2.3	7.7	45.0
Switzerland	2604	17.3	20.0	2.1	5.8	37.0
Taiwan	1664	15.0	16.8	2.0	4.9	35.2
Turkey	1022	15.6	21.2	2.2	5.4	30.4
United Kingdom	3438	17.5	20.9	2.3	5.6	35.2
United States	3419	17.9	16.6	2.2	5.8	40.8
Mean		18.3	21.0	2.2	6.2	38.9
BW SITC-5, 90-01	2731	5.6		1.6	2.7	9.7

$^{(a)}$ Mean elasticity, weighted by the total value of imports in good k during the whole period.
$^{(b),(c),(d)}$ Estimation period : 1995-2003 (b), 1994-2002 (c), 1996-2003 (d).

Table A.3: Descriptive statistics on the sectoral measurement errors[a], between 1994 and 2003

Importing country	Nb of sectors	including $sh6$	$sh4$	$sh2$	Weighted Mean[b]	P5	Median	P95
Argentina	503	361	71	71	0.95	0.28	1.02	5.92
Australia	1465	1253	167	45	0.90	0.29	0.95	2.13
Belgium[c]	3026	2853	151	22	0.96	0.54	0.98	1.53
Brazil	1117	919	131	67	0.88	0.28	0.93	2.29
Canada[d]	2773	2605	146	22	0.82	0.38	0.98	1.58
China	2530	2336	165	29	0.94	0.28	0.94	1.64
Denmark	1720	1530	150	40	0.93	0.44	0.99	2.15
France	3772	3652	110	10	0.96	0.60	0.98	1.47
Germany	3753	3633	112	8	0.98	0.63	0.99	1.46
Greece	1372	1205	118	49	0.97	0.33	0.95	1.73
Hong Kong	2737	2556	154	27	0.87	0.41	0.99	2.21
India	838	645	125	68	0.69	0.11	0.83	2.28
Indonesia	1301	1106	141	54	0.86	0.23	0.95	2.35
Italy	3177	3026	135	16	0.95	0.50	0.98	1.66
Japan	3188	3017	155	16	0.96	0.44	0.99	1.72
Korea	1944	1737	176	31	0.93	0.32	0.96	1.75
Mexico	3015	2833	159	23	0.90	0.35	0.96	1.47
Netherlands	3231	3062	150	19	1.00	0.49	0.99	1.79
Poland	1432	1250	132	50	0.86	0.31	0.92	1.93
Russia[e]	1815	1641	125	49	0.92	0.33	0.98	2.72
Singapore	1787	1592	159	36	0.85	0.34	0.97	2.10
Spain	3118	2969	126	23	0.94	0.45	0.97	1.58
Sweden	2080	1900	141	39	0.97	0.51	0.98	2.46
Switzerland	2692	2502	164	26	0.87	0.55	0.99	1.74
Taiwan	1794	1575	177	42	0.84	0.33	0.97	2.32
Turkey	1133	957	120	56	0.88	0.20	0.88	1.96
U.Kingdom	3574	3435	129	10	0.94	0.50	0.98	1.69
United States	3504	3395	95	14	0.91	0.43	0.98	1.33
Mean					0.91	0.39	0.96	2.03
BW SITC5, 90-01	1927					0.27	0.97	2.82

[a] The sectoral price measurement error appearing when omitting changes in the supply of variety available in a given market $(k; j)$ is computed as the ratio of λ_{kjt} over λ_{kjt-1} where $\lambda_{kjt} \equiv \sum_{i \in C_{kj}} Val_{kijt} / \sum_{i \in C_{kjt}} Val_{kijt}$ is the share of "common" varieties (i.e. varieties imported in both periods) in the nominal consumption of period t. This bias is computed on the whole 1994-2003 period, i.e. for $t-1 = 1994$ and $t = 2003$. It measures the total price effect of omitting changes in the supply of variety during this period.

[b] Geometric mean, weighted by the total value of sectorial imports in 2003.

[c],[d],[e] Estimation period: 1995-2003 (b), 1994-2002 (c), 1996-2003 (d).

Table A.4: Agregated price measurement bias on the 1994-2003

Importing Country	$CPI/EPI^{(a)}$ 94-03	Annual bias (%)	Annual bias (%) computed with a constant elasticity of substitution:			
			$\sigma = 2$	$\sigma = 2.66$	$\sigma = 5$	$\sigma = 8$
Argentina	1.000	0.00	0.06	0.04	0.02	0.01
Australia	1.013	0.15	1.15	0.69	0.29	0.16
Belgium	1.006	0.07	0.39	0.23	0.10	0.06
Brazil	1.074	0.80	0.97	0.58	0.24	0.14
Canada	1.027	0.33	1.28	0.77	0.32	0.18
China	1.013	0.15	0.18	0.11	0.05	0.03
Denmark	1.008	0.09	0.41	0.25	0.10	0.06
France	1.013	0.15	0.39	0.24	0.10	0.06
Germany	1.005	0.06	0.22	0.13	0.05	0.03
Greece	1.008	0.09	0.63	0.38	0.16	0.09
Hong Kong	1.038	0.41	0.75	0.45	0.19	0.11
India	1.042	0.46	2.36	1.41	0.58	0.33
Indonesia	1.019	0.20	0.97	0.58	0.24	0.14
Italy	1.008	0.09	0.40	0.24	0.10	0.06
Japan	1.006	0.07	0.31	0.18	0.08	0.04
Korea	1.006	0.07	0.46	0.28	0.11	0.07
Mexico	1.005	0.05	0.49	0.29	0.12	0.07
Netherlands	0.998	-0.02	-0.10	-0.06	-0.03	-0.01
Poland	1.029	0.32	1.19	0.72	0.30	0.17
Russian Federation	1.012	0.16	0.86	0.52	0.21	0.12
Singapore	1.061	0.66	1.64	0.98	0.41	0.23
Spain	1.012	0.13	0.43	0.26	0.11	0.06
Sweden	1.011	0.13	0.29	0.17	0.07	0.04
Switzerland	1.011	0.12	0.84	0.51	0.21	0.12
Taiwan	1.040	0.43	1.28	0.77	0.32	0.18
Turkey	1.022	0.24	1.37	0.82	0.34	0.19
United Kingdom	1.008	0.09	0.46	0.28	0.12	0.07
United States	1.018	0.20	0.78	0.47	0.19	0.11
Mean$^{(b)}$	0.982	0.20	0.73	0.44	0.18	0.10
BW,90-01	0.950	0.47				

$^{(a)}$ Ratio of the "exact" aggregate price index over the "conventional" price index (omitting changes in the supply of variety available in the considered country) in 2003 when 1994 is taken as the base year. Exceptions are Belgium (base year=1995), Russian Federation (base year=1996) and Canada (price index in 2002). The ratio of 1.013 obtained for Australia means that omitting changes in the supply of varieties leads to overestimate this country's import price by 1.3% on the whole 1994-2003 period. These figures are converted into an annual measurement error in the next column (called "Annual Bias") using the following formula: $Annual\ Bias_{jt} = 100 * \left(\frac{CPI}{EPI}^{1/Nb\ Years} - 1\right)$. Thus, in Australia, the annual price effect is .15% of the conventional price index.
$^{(b)}$ Mean results across countries: Geometric mean for the CPI/EPI bias, arithmetical mean for biases in %.

Table A.5: Measurement bias of relative import prices with respect to the United States (cumulated over the 1994-2003 period)

Importing Country	g_{RB} (%)	g_{CRMP} (%)	g_{ERMP} (%)
Argentina	-1.8	-14.8	-13.2
Australia	-0.5	-7.4	-6.9
Belgium	-1.2	14.9	16.3
Brazil	5.2	29.0	22.6
Canada	1.1	2.5	1.4
China	-0.5	.	.
Denmark	-1.0	5.9	7.0
France	-0.5	-10.3	-9.8
Germany	-1.3	-7.6	-6.4
Greece	-1.0	5.4	6.5
Hong Kong	1.9	-12.8	-14.4
India	2.3	12.3	9.8
Indonesia	0.0	.	.
Italy	-1.0	19.2	20.4
Japan	-1.2	-7.7	-6.6
Korea	-1.2	-13.7	-12.7
Mexico	-1.4	.	.
Netherlands	-2.0	3.8	5.9
Poland	1.0	4.8	3.8
Russian Federation	-0.4	.	.
Singapore	4.0	-7.9	-11.4
Spain	-0.7	2.1	2.8
Sweden	-0.7	11.7	12.5
Switzerland	-0.7	0.3	1.0
Taiwan	2.0	.	.
Turkey	0.3	-95.8	-95.8
United Kingdom	-1.0	-2.7	-1.7

For each importing country, the first column called g_{RB} gives the cumulated measurement error in its relative import price with respect to the United States, when computed in a "conventional" way. It corresponds to the ratio of the CPI/EPIs provided in the first column of the previous table, in percent of the US one: $g_{RB} = 100 * \left(\frac{CPI/EPI}{CPI^{US}/EPI^{US}} - 1\right)$. A negative figure means that omitting changes in the supply of variety available in the compared countries leads to underestimate the country's relative import price. To get an idea of the magnitude of this effect, the second column (g_{CRMP}) gives the observed growth of the "conventional" relative import price over the 1994-2003 period, obtained from the IMF's series of import unit value index (Source: *International Financial Statistics*): $g_{CRMP} = 100 * \left(\frac{CPI}{CPI^{US}} - 1\right)$. Last, the third column (g_{ERMP}) is the "Exact" growth of the country's import price index relative to the United States defined as: $g_{ERMP} = 100 * \left(\frac{1-g_{CRMP}}{1-g_{RB}} - 1\right)$.

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