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ESSAYS ON INTERNATIONAL PRICES AND THE SUBJACENT MARKET STRUCTURE

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Orientador: Renato Galvão Flôres Junior

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Agradeço,

Aos meus pais, pelo amor e apoio incondicional

Ao meu orientador, Renato Flôres, por acreditar e guiar este projeto a uma boa conclusão

Aos colegas da Divisão do Balanço de Pagamentos do Banco Central Do Brasil, pela amizade, encorajamento e auxílio técnico inestimável

Aos colegas da EPGE, que compartilharam comigo esta experiência

À minha esposa, Juliana, por tudo, por meu coração

Resumo

Resumo. Esta tese utiliza a informação contida em preços internacionais para identificar parâmetros de modelos de comércio sob competição imperfeita, desta forma permitindo inferência sobre o comportamento das exportações, sobre os ganhos de troca da abertura comercial e sobre a variedade de bens produzidos domesticamente. Em primeiro lugar, investigamos o repasse cambial, no longo prazo, para os preços praticados por exportadores brasileiros. O foco no longo prazo permite controlar os efeitos da rigidez de preço no curto prazo, de maneira que o repasse incompleto evidencie competição imperfeita com preços flexíveis. Em segundo lugar, calculamos os ganhos de troca de novas variedades de bens importados baseando-nos em estimativas para as elasticidades de substituição desagregadas. Finalmente, qualificamos a ênfase da literatura de comércio em ganhos de eficiência no lugar de ganhos de variedade, demonstrando que a variedade de bens produzidos domesticamente se amplia após aberturas comerciais desde que as firmas tenham uma margem de decisão em bens intermediários ou na qualificação da mão de obra.

Palavras-Chave: comércio internacional, competição imperfeita, preços internacionais, repasse cambial, ganhos de troca, variedade de bens

Abstract. The thesis uses international price data to identify parameters of trade models with imperfect competition, therefore allowing inference on exchange rate behavior, gains from trade and variety of domestic goods. First, we investigate Brazilian exporters pricing behavior over the long-run following destination specific exchange rate shocks. We find evidence of incomplete exchange-rate pass-through in the long-run, which supports the market structure explanations over short-run sticky-price explanations. Second, we calculate import price indexes and the implied welfare gains from new varieties of imported goods, based on disaggregated estimates of elasticity of substitution parameters. Finally, we qualify standard results in the literature that point to a reduction in domestic varieties after trade liberalization; domestic varieties may expand if we introduce an additional margin in firms' technology, such as intermediate goods or high skilled labor.

Keywords: international trade, imperfect competition, international prices, exchange-rate pass-through, gains from trade, variety of goods.

Contents

Introduction

This thesis is composed of three essays with a common interest on trade under imperfect competition. From a technical standpoint, we use different strategies to identify primitive parameters from information on international prices. From an applied standpoint, the primitive parameters underlying price behavior have wide-range implications; from the conduct of exchange-rate policy to the welfare consequences of trade; from aggregation biases in trade elasticities to the skill-premium consequences of foreign competition.

The first chapter investigates Brazilian exporters pricing behavior over the long-run following destination specific exchange rate shocks. The panel cointegration method of Bai, Kao and Ng (2009) is shown to identify the long-run parameter of interest. We find evidence of incomplete exchange-rate pass-through in the long-run, which supports the market structure explanations of Krugman (1986) over short-run sticky-price explanations. Optimal incomplete pass-through implies lower risk premium from short-run sticky-prices; therefore, it changes incentives for exchange rate policy.

The second chapter estimates import price indexes and the implied welfare gains from new varieties of imported goods. A lot of care is taken with heterogeneity in consumers' taste parameters over different kinds of goods, as well as heterogeneity in firms' technology parameters over different intermediate goods. To identify this wealth of elasticities of substitution parameters, we assembled import price data for a large array of goods and explored heteroscedasticity in the panel dimension of each of these data sets. Elasticities of substitution are shown to have systematic relations to market structure indicators. They also point to aggregation bias in standard estimates of aggregate elasticities of substitution and aggregate price-indexes.

The third chapter qualifies standard results in the literature of trade and heterogeneous firms that point to a reduction in domestic varieties after trade liberalization. We show domestic varieties may expand if we introduce an additional margin in firms' technology, such as intermediate goods or high skilled labor. Conditions on primitives and observable prices correlates allow inference on general equilibrium effects over domestic varieties production after trade opening.

Pricing-to-market by Brazilian Exporters: a Panel Cointegration Approach

Abstract

This paper investigates Brazilian exporters pricing behavior over the long-run following destination specific exchange rate shocks. The panel cointegration method of Bai, Kao and Ng (2009) is shown to identify the long-run parameter of interest. The method crucially depends on identification and controlling for the common trend in prices to different countries, a trend which is structurally interpreted, like originally proposed by Kneeter (1989), as the exporter's marginal cost. We find evidence of incomplete exchange-rate pass-through in the long-run, which supports the market structure explanations of Krugman (1986), known in the literature as pricing-to-market, over contending short-run sticky-price explanations. The degree of long-run pass-through is also shown to be positively related to technological intensity in the sector, a proxy for low elasticity of substitution of varieties.

JEL Classification: C52, D43, F12, F14,F31, L13

Keywords: exchange rate pass-through; panel cointegration, sticky-price, market structure, pricing-to-market.

1. Introduction

With sufficiently segmented international markets, exporters may set country specific prices to reflect local demand and competition conditions, a behavior Krugman (1986) called pricing-to-market. The concept is relevant for many empirical puzzles in international economics, such as incomplete exchange rate pass-through to exporter prices and persistent deviations from purchasing power parity. The main advantage over contending explanations, the leading one being sticky prices set by exporters in local currencies, is the optimal character of the pricing rules[1] and the possible accounting for persistent shocks on relative international prices[2]. This paper follows the empirical tradition of identifying microeconomic pricing-to-market behavior by resort to exporter's destination specific mark-up adjustments after exchange rate shocks [Krugman (1986), Kneeter (1989)]. Advancing previous methods, we look for credible estimates of long run pricing-to-market effects on Brazilian exports. By definition, the effect cannot be attributed to sticky price and other short run explanations of incomplete exchange-rate pass-through. On the contrary, it amounts to strong evidence in favor of market structure explanations, the more so considering it refers to exporters from a developing country where it is least expected. Additionally, we look for patterns of behavior across different industries, a significant undertaking in face of the large number of imperfect competition models that may explain the results.

Krugman's (1986) original strategy to identify pricing-to-market behavior was trying to control for exogenous trends in prices common across all export destinations in order to recognize divergent trends after a large real exchange rate shock; this was implemented for German export data. Kneeter (1989) developed this basic insight further. The author proposed a panel framework that controls for common trends in prices by the inclusion of a time effect. He also provided structural interpretation for the common trend as the production cost of the marginal unit, such as can be deduced from the optimization

[1] Sticky local currency prices have welfare implications under flexible exchange rate regimes through the addition of a risk premium term to import and export prices [Sutherland (2005)]; by reducing optimal variability, pricing-to-market have a direct bearing on risk premium and welfare.

[2] Atkeson and Burstein (2008) and Ravn (2001) have established the quantitative importance of general equilibrium models with pricing-to-market, as opposed to sticky prices, in reproducing broad features of international relative prices. While these authors focus on segmentation of producer prices on home and foreign markets, we look at segmentation of export price across foreign markets.

problem of a representative firm exporting to many destinations. The model proposed in this paper has very similar interpretation; so, it is worth getting at some details of Kneeter's econometric specification. The author singles out the bilateral exchange rate, measured as the price of exporter currency in each destination country's currency, as the relevant country specific shock, from which to gauge the presence and quantitative magnitude of pricing-to-market behavior. Measuring prices in the exporter's currency, the degree of pricing-to-market can be assessed by the partial effect of the bilateral exchange rate on export price while holding marginal cost fixed. Expressing the variables in logarithms, this coefficient measures the effect of the exchange rate on exporter's mark-up over marginal costs. Also, since foreign prices are just the sum of the exporter price and the exchange rate, one plus the coefficient measures the exchange rate pass-through to foreign consumers due to mark-up adjustments.

An important result from Kneeter (1989) is that U.S. exporters in most industries appear to fully or over pass-through exchange rate movements to foreign consumers. As noted by the author, instead of the usual assumption of foreign demand curve becoming more elastic as price rises, such coefficients would require just the opposite assumption (for mark-ups are inversely related to the elasticity of demand). This sort of result persisted throughout the literature: there is wide dispersion, usually around full pass-through levels and many coefficients have significant but counterintuitive signs. For a recent example, Méjan (2004) found exactly those results for massive data sets on Germany, United States, France, Italy, Japan and United Kingdom exporters, with volume and price data organized in much disaggregated sectors. On a more optimistic tone, Méjan interpreted the wide dispersion as room for structural, microeconomic explanations. In the most recently published paper on the subject, Bugamelli and Tedeschi (2008) report the same pattern of result after many variations of Kneeter's basic specification. Because their estimated coefficients bundle many different products, the dispersion is not as accentuated as in Méjan's paper. Bugamelli and Tedeschi experimented with product classifications in search for some microeconomic rationale for the resulting heterogeneity. They found some evidence of stronger pass-through for products characterized by increasing returns to scale or intense use of science, features often associated with oligopolistic industries. Most of the studies from the literature use annual data to dismiss dynamic considerations. There are

attempts at dynamic panel models with quarterly or monthly data, such as Takagi and Yoshida (2001), on Japanese exports. But the inclusion of lagged price as an explanatory variable, with the associated dynamic panel methodology, leads to the same pattern of results.

The main contribution from this paper is to extend Kneeter's panel method to allow for long-run relations between the variables and to actually estimate the long-run parameters. We trusts long run relations will provide a clearer picture on pricing-to-market behavior, with coefficients less dispersed, more plausibly signed and easier to connect with microeconomic fundamentals. In addition, because nominal rigidities have no place in the long-run, the estimated coefficients may discriminate between market structure and sticky prices explanations of incomplete pass-through. At a conceptual level, an important advance is to model the common marginal cost trend as a stochastic process on par with the price and exchange rate processes, with the explicit possibility of equilibrium relations among all of the variables. Long-run relations are modeled as cointegration among integrated processes, which is a restricted but useful interpretation. For example, this rules out mean reverting residuals with long memory, as well as overlooks possible inference problems from near unit root series. In a sense, though, the econometric method proposed in this paper is genuinely more general then previous methods used in the literature, since estimates are consistent even if some of the variables are stationary in levels as previous authors have maintained. The unit-root hypothesis is necessary only for a long-run interpretation of the coefficients; even so, the paper tests this null against stationarity with panel techniques that have greater power then single series tests.

The appropriate econometric theory to address long-run issues in a panel framework was only recently developed. Philips and Moon (1996) were the first to propose consistent estimators for panel cointegration vectors with the concomitant development of the asymptotic theory for sequential and simultaneous limits in the two panel dimensions. An important shortcoming was the assumption of independent errors along the cross-section dimension. To overcome this difficulty, Bai, Kao and Ng (2009) developed a "second-generation" framework where common factors, possibly with unit roots, capture the cross-sectional dependence. Since our structural model postulates the existence of cross-section dependence due to the common marginal cost to all export destinations, this seems as an

appropriate estimator. Indeed, the behavioral model proposed in this paper has an exact mapping to Bai, Kao and Ng econometric specification. Nevertheless, it should be stressed that, while the common factors are just an econometric device in their model, here we adopt the much stronger structural interpretation of a common marginal cost trend in export prices to different destinations[3]. As already mentioned, the estimator does not require pre-testing the variables for unit-root against the stationarity alternative. However, in order to check if our initial concerns were of any consequence, we also applied Bai and Kao (2002) panel unit root test to the bilateral exchange rate and export price series.

Another contribution from the paper is the use of product classification to uncover possible microeconomic patterns in the estimated pass-through coefficients. Although the approach is similar to Bugamelli and Tedeschi (2008), two important differences have a bearing on the results. First, the authors used unit values throughout, estimating the common effect for a group of products by including an associated dummy variable. In contrast, this paper bundles the products from the start and calculates price indexes on which the whole analysis is conducted. Second, the actual classification scheme is different; the one used here is closely based on technological intensity. As a research principle, the discipline of building the indexes before obtaining the estimates lends more credibility to any pattern eventually found, since one minimizes snooping through many different aggregations. It also potentially ameliorates the measurement error from using unit values as proxies for prices, under the assumption of independent errors. As for the classification system, we have used it in parallel research which actually suggests some connections with structural preference and market structure parameters.

The remainder of the paper is organized as follows. Section 2 describes the data and limitations. Section 3 reports results on panel unit-root tests relevant for interpreting the results. Section 4 is the heart of the paper, explaining the structural model, the econometric model and the estimation results. Section 5 discusses the panel cointegration results in relation to previous estimates and looks for microeconomic patterns. Section 6 concludes.

[3] Hatemi-j and Irandoust (2004) study long-run exchange rate pass-through to Swedish import prices in a cointegrated panel framework; the authors mention "pricing-to-market" although all price data refers to Sweden and have therefore no bearing on segmentation issues; the authors do not admit any structural common factor along the cross-section dimension and use first-generation panel estimators.

2. Data description and limitations

The data sample ranges from the first quarter of 1997 to the third quarter of 2006, which amounts to 39 periods. Depending on the sector, the number of export destination countries can be as little as 29 and as much as 53. On the one hand, the small sample size in the two panel dimensions, time and country, could raise problems for the asymptotic inferences. On the other hand, the estimator used here was shown to have good finite sample properties in simulation experiments reported by Bai, Kao and Ng (2006). Indeed, the mean bias and standard deviation of the estimates keep their good asymptotic properties in samples as small as 20 in both panel dimensions. More importantly, the estimator has much better properties in small samples than the alternatives[4].

As shown in Table 1, Brazilian exports of manufactures were classified in thirteen sectors according to technological intensity. There are many reasons to adopt some level of aggregation. First, this reduces data to manageable proportions. Second, price indexes average out measurement error in the price data. Third, it permits to investigate a possible relationship between pass-through and technological intensity, throwing some light on the microeconomic structure driving the results. Finally, Brazilian official institutions often use this classification scheme, which was developed by OCDE in 1995 (Classification of High-technology Products and Industries), thus facilitating research communication. Some sectors from the original classification, namely aviation, ship building and oil, have been excluded due to insufficient number of export destination countries. The sector share in the table refers to the value exported in the third quarter of 2006 relative to total manufacturing export.

Table 1 also reports the number of countries in each sector. The criteria for including a country in a particular sector were positive export to this country in all periods and for a significant fraction of the products from the sector. The selection rule ensures a balanced panel in each sector, as well as high quality export price indexes. The rule could lead to selection bias. However, the most likely reason for low trade volume is trade barrier,

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[4] See the Monte Carlo section in Bai, Kai and Ng (2006), with particular attention to table 1 and 2 and the column referring to the fully modified continuously updated estimator. Another justification for relying in such small samples is the freshness of the method and the possible implications for an important field of applied research.

and the pricing behavior exporters would have in case of no barrier should be independent of the actual level of the barrier. Additionally, excluded countries get none or just a small share of Brazilian exports, and it seems justified to give much less weight to these countries in the estimation of the pooled coefficient. The last section of the paper discusses the selection issue further.

Table 1. Industrial sectors by technological intensity: share and number of countries

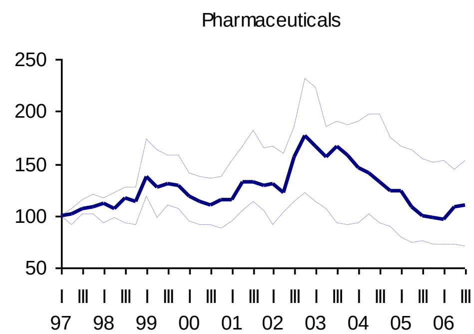
Technology	Industrial Sector	Share (%)	Countries
High	Pharmaceuticals	0,7	29
	Medical, precision and optical instruments	0,5	32
Medium-high	Eletrical machinery and apparatus	2,3	39
	Machinery and equipment	7,1	51
	Chemicals excluding pharmaceuticals	5,8	43
	Motor vehicles, trailers and semi-trailers	11,8	41
Medium-low	Rubber and plastics products	1,8	53
	Non-metallic mineral products	1,8	53
	Basic metals and fabricated metal products	12,8	34
Low	Food products, beverages and tobacco	17,9	46
	Wood, pulp, paper, paper products	5,4	48
	Other Manufacturing	1,0	33
	Textiles, leather and footwear	4,8	52
		73,7	29-53

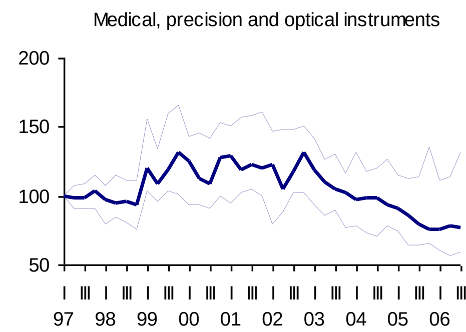
Export price indexes were calculated with unit values at the finest level of the Harmonized System, an international standard for commodity classification. The Fisher index formula was used after a preliminary trimming of too extreme variation, very unlikely to reflect price developments. For each country, in each sector, an export price series was constructed. Unit values are poor measures of price. Still, they are readily available and very much used in empirical studies on international prices. The inclusion of country specific effects should capture any measurement error that survives aggregation. As already mentioned, one reason to use aggregate price indexes is to average out the likely measurement errors in prices. On the other hand, the more aggregated sector classification we use, the less reasonable our structural interpretation of the data. Relative to other studies with a similar panel approach, our choice in this trade-off involves more aggregation.

Figure 1. Export price indexes by industrial sector; high and medium high technology

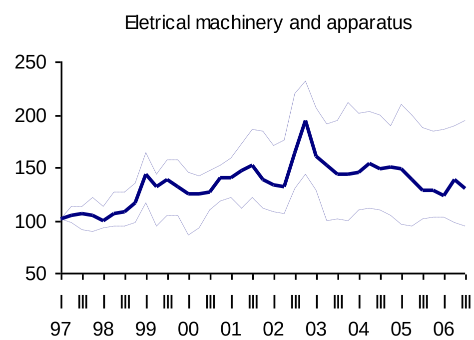
Median in bold-line; other quartiles in dashed-lines; 1997:I = 100; Monetary Unit = R$

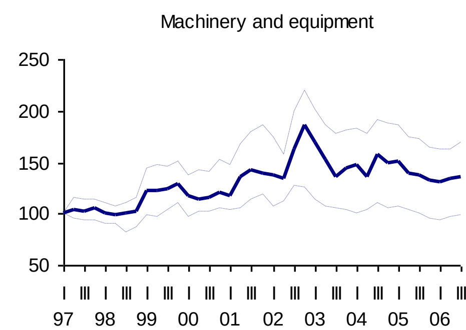
High Technology

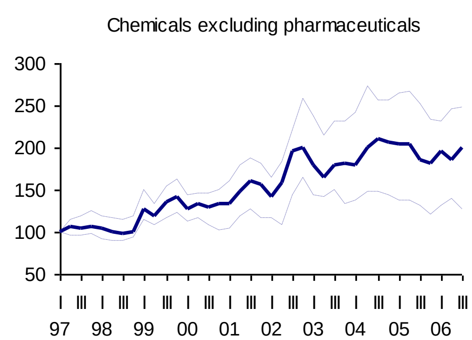




Medium-high Technology







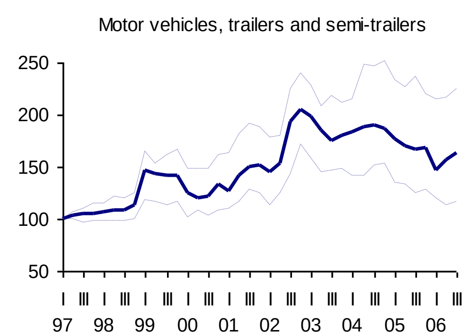
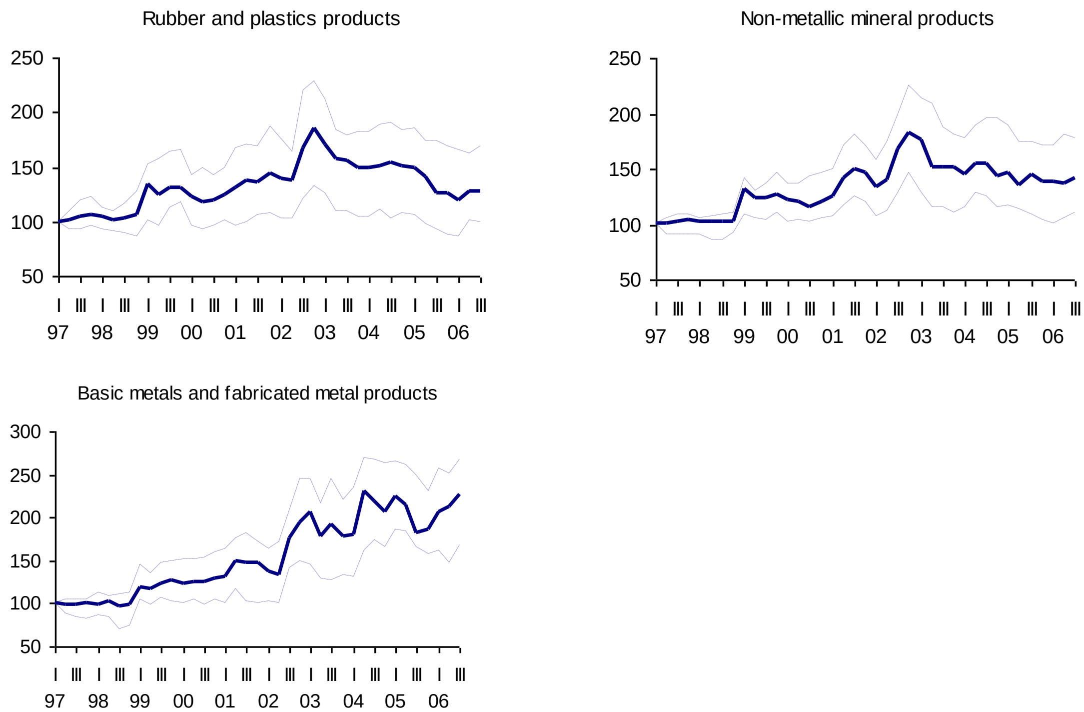


Figure 2. Export price indexes by industrial sector; medium-low and low technology

Median in bold-line; other quartiles in dashed-lines; 1997:I = 100; Monetary Unit = R$

Medium-low Technology



Low Technology

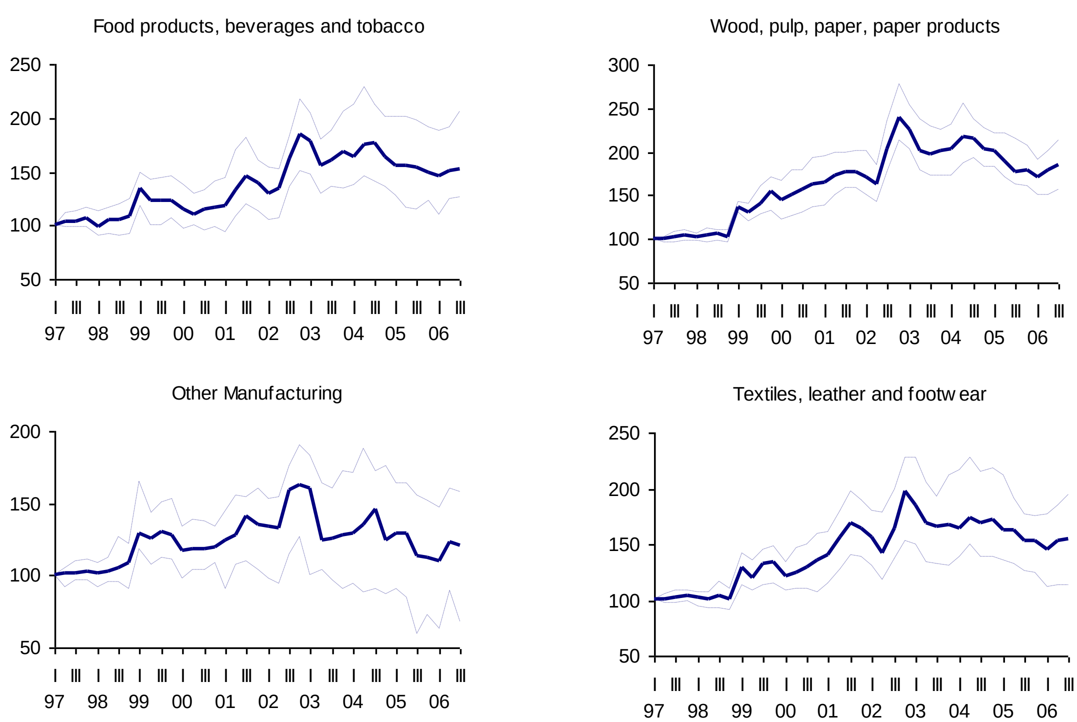


Figure 3. Real bilateral exchange rate

Median in bold-line; other quartiles in dashed-lines; 1997:I = 100

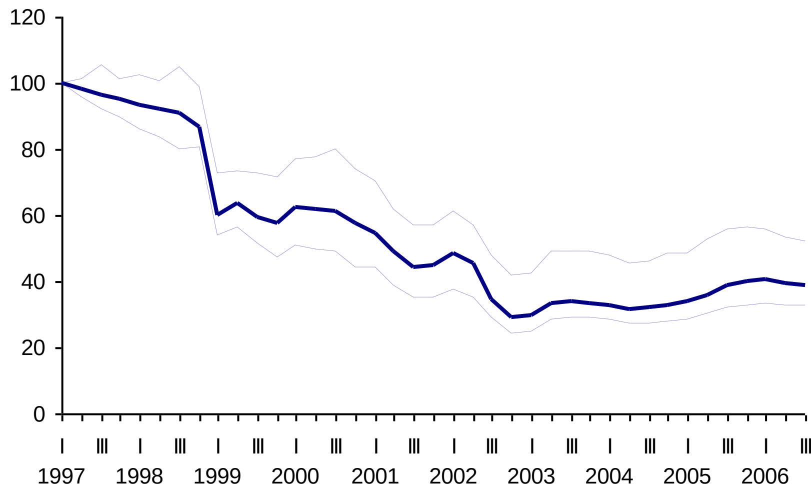


Figure 1 and 2 summarize the export price data for each sector. The median is a rough estimate of a common trend component, while the other quartiles capture how closely country data follow this trend. As can be seen by the widening distance of the quartile-band in most sectors, the common trend appears to have a weak influence in the price data. This will be more formally studied in the next section.

Bilateral exchange rate series were corrected for consumer price inflation in each export destination country. Both the nominal exchange rate and consumer price data were obtained from the International Financial Statistics (IFS) database. After the correction, we obtain a 'real' exchange rate measure, from the point of view of consumers in each export destination country. Indeed, we will often refer to the corrected exchange rate as the bilateral real exchange rate. One could argue that sector specific price inflation in each foreign country should be used in place of consumer price inflation, but this kind of data is not easily obtainable. Figure 3 summarizes exchange rate data. Compared to the price series, the common trend captured by the median seems to be much more important for exchange rates, with all the country series following it very closely.

3. Results on panel unit-root

Before estimating the parameter of interest for each industry, we test all the series for unit-root using a panel technique developed by Bai and Kao (2004). The procedure is different from the usual time-series ones in at least two respects; first, it can spot low frequency movements that would otherwise go undetected, as long as they are shared by the series in the panel; second, it can pool the unit-root tests, making it harder to incorrectly accept the unit-root null for a group of time-series. The first feature, which is unique to Bai and Kao approach, will be particularly important for correctly classifying the export price index series. It is important to highlight that none of this pre-testing is necessary for the estimation of the pricing-to-market coefficient; the estimator used in the next section would be consistent even if some of the series were stationary (but it would converge more slowly). Still, pre-testing is crucial for interpreting correctly the estimated parameters.

The test decomposes each series from a panel into a common factor and an idiosyncratic component. The estimation procedure makes no assumption about integration order of the components, thus allowing testing each component separately. Common factors serve as an econometric device to capture cross-section dependence, thus overcoming the most serious deficiency of first-generation panel tests for unit-root. In contrast to the model in the next section where the structural dependency between variables is explicitly addressed, in this section we model each variable in separation from the other.

Formally, each time-series from a given panel of series is decomposed as:

$$y_{it} = d_{it} + \lambda_i' F_t + \nu_{it} \qquad (1)$$

where $i=1..N$ indexes the panel unit, $t=1..T$ indexes the time periods, d_{it} is a determinist trend, F_t is a vector of common factors, λ_i a vector of factor loadings, ν_{it} is the error component and y_{it} is the variable of interest. The common factors are estimated as principal components from the differenced series summed up to each period. The number of factors to be included in the model for each panel was selected by information criteria as suggested by Bai and Ng (2002).

For the export price series, each sector was treated as a separate panel data set, and each of these panels was modeled as in equation (1), with *i* indexing the export destination country. This is a reasonable setup given that the common factors influencing the price to each destination country is most likely sector specific; we have just imposed this by assumption. Admitting at most three factors, a single factor was selected in all panels.

For the exchange rate data, the common factors are mostly associated with macroeconomic conditions unrelated to the sectors. Thus, all the export destination countries were pooled in a single panel, as in equation (1), with *i* indexing the country. Again, admitting at most three factors, a single factor was selected for the exchange rate.

Since there are many panels, each with several time-series, we report only summary information on the testing results. Qualitative results are not sensitive to the specification of the deterministic component; but we report results under different assumptions.

Table 2. PANIC test for unit root; one factor; constant and time trend included *39 time periods*

Export price by technology and sector		Countries	Explained by Factor >25%	Explained by Factor >40%	Series %Rej	Factor t	Factor p	Error %Rej	Error Pool
High	Pharmaceuticals	29	34%	10%	10%	-2,86	0	28%	>
	Medical, precision and optical instruments	32	16%	3%	18%	-1,63	1	34%	>
Medium-high	Eletrical machinery and apparatus	39	8%	3%	20%	-4,59	1	26%	>
	Machinery and equipment	51	4%	2%	27%	-3,28	0	37%	>
	Chemicals excluding pharmaceuticals	43	53%	28%	23%	-3,00	0	40%	>
	Motor vehicles, trailers and semi-trailers	41	24%	5%	20%	-2,02	2	37%	>
Medium-low	Rubber and plastics products	53	4%	2%	15%	-4,52	1	25%	>
	Non-metallic mineral products	53	43%	17%	19%	-3,43	0	30%	>
	Basic metals and fabricated metal products	34	18%	6%	35%	-2,94	3	47%	>
Low	Food products, beverages and tobacco	46	13%	4%	15%	-2,61	2	24%	>
	Wood, pulp, paper, paper products	48	73%	46%	8%	-2,14	0	19%	=
	Other Manufacturing	33	6%	6%	27%	-2,42	1	39%	=
	Textiles, leather and footwear	52	50%	29%	17%	-2,64	0	31%	>

Bilateral exchange rate corrected for consumer price	n	Explained by Factor >75%	Explained by Factor >90%	Series %Rej	Factor t	Factor p	Error %Rej	Error Pool
	68	76%	44%	0%	-1,07	0	11%	=

Notes: (i) "Explained by Factor" is the fraction of n countries where the common factor explains at least 25, 40, 75 or 90 percent of the variation. (ii) "%Rej" is the fraction of the original or error series for which the unit-root hypothesis is rejected. (iii) The Dickey-Fuller t-test statistic and lag order are shown only for the common factor, and rejection is indicated be an underline. (iv) A greater

Table 2 reports the results when country specific intercept and time-trend are both included. The column entitled "Explained by Factor" reports the fraction of countries for which the common factor explains at least 25 or 40 percent of the variation in the original series, for the case of export price indexes, and at least 75 or 90 percent of the variation, for the case of the exchange rate series. These bands were selected to highlight where most of the action is occurring. Confirming the informal impression from the last section, the common factor accounts for very little of the variation in the export price series in all sectors, but is the main driving force of exchange rates.

The last main columns of Table 2 refer to Augmented Dickey-Fuller t-tests applied respectively to the original series, to the estimated common factor and to the estimated error. The test for the error series does not include a constant and neither a time trend (the critical value is -1.94). Both deterministic components are included in the tests for the original and the factor series (the critical value is -3.41). The number of lags was selected by sequential F-tests, with a liberal 50% significance level, admitting at most three lags.

In the case of the original series and the error series, only the fraction of rejections of the unit root hypothesis is reported. Thus, the percentage indicates the fraction of countries with stationary series. In the case of the common factor series, the table reports the test statistic and the number of selected lags; rejection is indicated in boldface.

Finally, the last column is the result of a pooled test for the unit-root null on the residual series (tests can be pooled on the residuals because all common factors where already extracted). Only series regarded as non-stationary according to the single-series test where included in the pool. Therefore, the idea is to determine if at least one of this series is actually stationary. Accordingly, the “greater than” sign indicates when the fraction of stationary series is likely larger than indicated in the previous column.

Analyzing the lower part of Table 2, we conclude there is strong evidence of unit-root for all the exchange rate series. In fact, the panel testing procedure reinforces results obtained for the original time series. The main source of non-stationarity is the common factor, although most of the idiosyncratic components also display a unit-root.

Results are more interesting for the export price indexes, as displayed in the upper part of Table 2. One of the motivations for Bai and Kao (2004) was the possibility of a very

small non-stationary component, which would be hard to capture by traditional tests. This appears to be exactly the case for the export price series. While the common factor explains little of the variation, they have a unit-root in most sectors. In comparison, the idiosyncratic components are often stationary and explain a lot of the variation in the original series, making it hard to detect a unit-root in the series. Indeed, as can be seen from the fifth column from Table 1, single-series tests often reject a unit-root, completely missing the presence of the low frequency common component. In the three sectors where the common factor appears to be stationary, the idiosyncratic components have unit-roots for 70% or more of the destination countries.

Table 3. PANIC test for unit root; only constant included *39 time periods*

			Explained by Factor		Series	Factor		Error	
Export price by technology and sector		Countries	>25%	>40%	%Rej	t	p	%Rej	Pool
High	Pharmaceuticals	29	34%	10%	0%	-2,87	0	3%	>
	Medical, precision and optical instruments	32	16%	3%	13%	-1,18	1	16%	>
Medium-high	Eletrical machinery and apparatus	39	8%	3%	15%	$\underline{-3,98}$	1	15%	>
	Machinery and equipment	51	4%	2%	22%	-1,39	3	10%	>
	Chemicals excluding pharmaceuticals	43	53%	28%	9%	-1,87	0	14%	>
	Motor vehicles, trailers and semi-trailers	41	24%	5%	17%	-3,55	0	7%	>
Medium-low	Rubber and plastics products	53	4%	2%	8%	$\underline{-4,52}$	1	8%	>
	Non-metallic mineral products	53	47%	21%	19%	$\underline{-2,24}$	2	8%	>
	Basic metals and fabricated metal products	34	35%	6%	6%	-1,66	1	9%	>
Low	Food products, beverages and tobacco	46	13%	4%	11%	-2,19	2	9%	>
	Wood, pulp, paper, paper products	48	73%	50%	2%	-1,32	0	6%	>
	Other Manufacturing	33	6%	6%	18%	-1,88	1	21%	=
	Textiles, leather and footwear	52	52%	29%	2%	-2,63	0	6%	>
			Explained by Factor		Series	Factor		Error	
Bilateral exchange rate corrected for consumer price		n	>75%	>90%	%Rej	t	p	%Rej	Pool
		68	94%	73%	4%	-1,96	0	10%	=

Notes: (i) "Explained by Factor" is the fraction of n countries where the common factor explains at least 25, 40, 75 or 90 percent of the variation. (ii) "%Rej" is the fraction of the original or error series for which the unit-root hypothesis is rejected. (iii) The Dickey-Fuller t-test statistic and lag order are shown only for the common factor, and rejection is indicated be an underline. (iv) A greater

Table 3 shows results without the time trend. The general picture is much like before, with the exception of a larger proportion of non-stationary idiosyncratic components. There is also disagreement on the integration order of common factors for two

of the industrial sectors. But common factors still account for very little of the variation in the export price series, the opposite being the case for the exchange rate series.

4. Results on panel cointegration

We present coefficient estimates for the partial effect of the bilateral exchange rate on export price controlling for marginal cost. The most direct way of attributing meaning to this coefficient is through the first order condition for the profit maximization problem of a representative domestic firm from a given industry exporting to several destinations. The bilateral exchange rate enters the first-order condition of the respective destination, and may be interpreted as a demand or cost shift. Comparative statics imply mark-up adjustments specific to each market. Equilibrium effects may be incorporated implicitly in residual demand curves or explicitly in additional first order conditions. The econometric model proposed here is a log-linear approximation of the bilateral exchange rate comparative statics effect which implicitly assumes small higher order effects.

There are many ways to model this formally. The essential point, though, is variable demand elasticity along the demand curves. Kneeter (1989) presents a simple model for an arbitrary residual demand function. Dornbush (1987) explores the residual demands that emerge from different market structures. Atkeson and Burstein (2008) take demand elasticity as a function of market share and preferences for variety. The structural interpretation is essentially static, but the inclusion of a time trend will hopefully capture changes in equilibrium market structure. The next section discusses structural models in greater detail to anchor theoretical expectations on microeconomic fundamentals.

The econometric model for the panel data set for an arbitrary sector is,

$$p_{it} = d_{it} + \beta e_{it} + \lambda_i \, MgC_t + u_{it} \tag{2}$$

where $i=1..N$ indexes the country, $t=1..T$ indexes the time periods, d_{it} is a determinist term, p_{it} is the export price in domestic currency, e_{it} is the real bilateral exchange rate (real price of domestic currency in country i), MgC_t is the common marginal cost trend and u_{it} is an error term. Except for the error term, all random variables are assumed integrated of

order one; this is in conformity with results from the last section. The error term u_{it} may be at most weakly dependent on the time and the cross-section dimension.

The coefficient β measures the partial effect of the bilateral real exchange rate, and is the parameter we are interested at. This coefficient is pooled across destination countries. As a consequence, in case there is significant heterogeneity of firm behavior across markets, the parameter actually represents an average effect, as noted by Philips and Moon (1996). The coefficient λ_i measures the partial effect of the marginal cost. It is country specific, which allows for different compositions of export bundles to different countries, or any other reason for some degree of freedom in the assessment of marginal cost relevant for a destination country.

The model follows very closely Bai, Kao and Ng (2009), with the marginal cost trend in equation (2) standing for a common factor in their terminology. Given our desired structural interpretation, we did not use information criterion to set the number of common factors, just fixing it at one as in equation (2). Except for low-technology goods, we argue below results are not very sensitive to this assumption. As for the deterministic term, we have tried the model with and without a deterministic trend. The qualitative results are reasonably close to each other, but with the trend better adhering to the data. The estimation method once again uses principal components to extract the common factor implicit in the export price indexes. Since this allows one to recover the marginal cost trend only up to a linear transformation, there is no expected sign pattern on λ_i's coefficients. For this reason, we will not emphasize these estimates, rather concentrating on the exchange rate effect.

As a matter of comparison, we also estimated two alternative models. The first is Kneeter's specification, for which variables enter in first-difference and the marginal cost trend enter as a fixed time effect (differencing is necessary, given our results on unit-roots from the previous section). The second is Philips and Moon (1996) fully modified estimator, which amounts to model (2) with λ_i's restricted to be zero; that is, disregarding any cross-sectional dependency.

The first results are presented on Table 4. They were obtained under the assumption of both a constant and a time trend present in the country specific determinist term. To obtain the degree of exchange rate pass-through to foreign consumers, just add one to the exchange rate coefficients on the table.

Table 4. Panel cointegration results; one factor; constant and time-trend included *39 time periods*

			Alternative estimates			Final estimate (β)			
		Countries	ols	fm-ols	cup	fm-cup	fit	e/cmg	%Rej
High	Pharmaceuticals	29	-0.19	-0.53	-0.44	-0.4	54%	1.31	69%
						(0.038)			
	Medical, precision and optical instruments	32	-0.1	-0.49	-0.17	-0.26	45%	0.26	63%
						(0.05)			
Medium-high	Eletrical machinery and apparatus	39	-0.1	-0.47	-0.24	-0.23	45%	0.4	54%
						(0.046)			
	Machinery and equipment	51	-0.18	-0.44	-0.4	-0.38	31%	2.72	61%
						(0.024)			
	Chemicals excluding pharmaceuticals	43	-0.05	-0.35	-0.45	-0.46	38%	7.23	60%
						(0.026)			
	Motor vehicles, trailers and semi-trailers	41	-0.08	-0.39	-0.53	-0.53	55%	3.69	68%
						(0.026)			
Medium-low	Rubber and plastics products	53	-0.08	-0.38	-0.52	-0.46	52%	2.21	55%
						(0.03)			
	Non-metallic mineral products	53	-0.11	-0.4	-0.42	-0.4	49%	3.1	66%
						(0.022)			
	Basic metals and fabricated metal products	34	0.03	-0.25	-0.41	-0.53	45%	2.92	56%
						(0.039)			
Low	Food products, beverages and tobacco	46	-0.1	-0.28	-0.5	-0.46	44%	2.55	61%
						(0.024)			
	Wood, pulp, paper, paper products	48	-0.07	-0.58	-0.56	-0.59	80%	3.07	56%
						(0.024)			
	Other Manufacturing	33	-0.14	-0.41	-0.41	-0.38	36%	4.77	64%
						(0.033)			
	Textiles, leather and footwear	52	0.02	-0.45	-0.45	-0.43	47%	2.81	69%
						(0.02)			

Notes: (i) "ols" is the estimator from Kneeter. (ii) "fm-ols" is the estimator from Philips and Moon. (iii) "cup" is the continously updated estimator from Bai, Kai and Ng. (iv) "cup-fm" is the fully modified version of the last estimator, with standard error in parentheses. (v) The fit is the median explained variance among destination countries. (vi) The "e/cmg" is the median ratio of variances of the exchange rate effect and the marginal cost effect. (vii) "%Rej" is the fraction of rejections of the unit-root null aplied to model residuals of each country in the sector.

Like the previous empirical literature, Kneeter's specification leads to very high levels of pass-through and to a few counterintuitive signs. The Philips and Moon specification appears to be more sensible, but it does not account for cross-sectional dependency (which is very likely given the results from the previous section), and it does not attempt to control for marginal cost, which is necessary for an economic interpretation.

The estimator from Bai, Kao and Ng addresses both issues simultaneously. Their fully modified continuously updated estimator, which has a normal asymptotic distribution, is reported on the fifth column, with standard errors in the column next to it. All coefficients are negative at any reasonable significance level. The degree of pass-through implied by the coefficients is much more plausible then the one implied by previous methods. Comparing the fifth and the third column, we see that cross-sectional dependence is an important issue for most of the sectors. Now comparing the fifth with the fourth

column, we notice that endogeneity of regressors, probably of the common marginal cost trend, may be an issue in some industrial sectors. Given the superior performance of the fully modified continuously updated estimator in the Monte Carlo experiments conducted by Bai, Kao and Ng (2009), we take this as our final estimator, and all the other summary measures in Table 4 take this reference point.

Table 5. Panel cointegration results; one factor; only constant included *39 time periods*

			Alternative estimates			Final estimate (β)			
		Countries	ols	fm-ols	cup	fm-cup	fit	e/cmg	%Rej
High	Pharmaceuticals	29	-0.19	-0.13	-0.56	-0.56	56%	1.28	21%
						(0.045)			
	Medical, precision and optical instruments	32	-0.09	-0.05	-0.33	-0.32	53%	0.53	47%
						(0.041)			
Medium-high	Eletrical machinery and apparatus	39	-0.11	-0.27	-0.42	-0.42	52%	1.9	41%
						(0.033)			
	Machinery and equipment	51	-0.18	-0.3	-0.49	-0.5	55%	3.54	37%
						(0.03)			
	Chemicals excluding pharmaceuticals	43	-0.06	-0.51	-0.39	-0.4	74%	1.04	47%
						(0.033)			
	Motor vehicles, trailers and semi-trailers	41	-0.08	-0.44	-0.48	-0.48	71%	2.51	34%
						(0.039)			
Medium-low	Rubber and plastics products	53	-0.09	-0.24	-0.42	-0.38	55%	3.06	36%
						(0.03)			
	Non-metallic mineral products	53	-0.12	-0.36	-0.43	-0.38	63%	2.88	36%
						(0.033)			
	Basic metals and fabricated metal products	34	0.02	-0.54	-0.07	-0.07	78%	0.01	35%
						(0.064)			
Low	Food products, beverages and tobacco	46	-0.1	-0.36	-0.52	-0.46	74%	2.75	39%
						(0.032)			
	Wood, pulp, paper, paper products	48	-0.07	-0.61	-0.1	-0.09	88%	0.03	29%
						(0.038)			
	Other Manufacturing	33	-0.14	-0.21	-0.41	-0.36	60%	2.13	55%
						(0.038)			
	Textiles, leather and footwear	52	0.02	-0.37	-0.46	-0.47	71%	5.06	48%
						(0.022)			

Notes: (i) "ols" is the estimator from Kneeter. (ii) "fm-ols" is the estimator from Philips and Moon. (iii) "cup" is the continously updated estimator from Bai, Kai and Ng. (iv) "cup-fm" is the fully modified version of the last estimator, with standard error in parentheses. (v) The fit is the median explained variance among destination countries. (vi) The "e/cmg" is the median ratio of variances of the exchange rate effect and the marginal cost effect. (vii) "%Rej" is the fraction of rejections of the unit-root null aplied to model residuals of each country in the sector.

The “fit” column on Table 4 refers to the median explained variance, where the median is taken with respect to export destination countries. The simple model with exchange rates and a single common factor as regressors explains about 50% of the variation in the export price data. Most of the explained variation comes from the exchange rate effects. This can be read from the next to last column, where it is shown the median

ratio of the variation due to exchange rate and the variation due to the common factor. The last column presents the fraction of countries residual series where the unit-root null is rejected. Critical values were obtained admitting three cointegrated variables. Given the low power of single-series tests, there is very strong evidence that residuals are stationary. This means the relationship summarized by the coefficient is not spurious.

Table 6. Sensitivity of parameter estimate (β) to the number of factors

			Number of Factors		
		Countries	1	2	3
High	Pharmaceuticals	29	-0.4	-0.57	-0.55
			(0.038)	(0.035)	(0.033)
	Medical, precision and optical instruments	32	-0.26	-0.3	-0.47
			(0.05)	(0.045)	(0.03)
Medium-high	Eletrical machinery and apparatus	39	-0.23	-0.25	-0.31
			(0.046)	(0.041)	(0.035)
	Machinery and equipment	51	-0.38	-0.42	-0.55
			(0.024)	(0.025)	(0.022)
	Chemicals excluding pharmaceuticals	43	-0.46	-0.44	-0.54
			(0.026)	(0.024)	(0.026)
	Motor vehicles, trailers and semi-trailers	41	-0.53	-0.47	-0.43
			(0.026)	(0.022)	(0.026)
Medium-low	Rubber and plastics products	53	-0.46	-0.49	-0.47
			(0.03)	(0.026)	(0.024)
	Non-metallic mineral products	53	-0.4	-0.48	-0.51
			(0.022)	(0.022)	(0.018)
	Basic metals and fabricated metal products	34	-0.53	-0.53	-0.65
			(0.039)	(0.031)	(0.037)
Low	Food products, beverages and tobacco	46	-0.46	-0.46	-0.47
			(0.024)	(0.025)	(0.02)
	Wood, pulp, paper, paper products	48	-0.59	-0.12	-0.08
			(0.024)	(0.032)	(0.032)
	Other Manufacturing	33	-0.38	-0.38	-0.3
			(0.033)	(0.027)	(0.025)
	Textiles, leather and footwear	52	-0.43	-0.49	0.08
			(0.02)	(0.02)	(0.034)

Note: Fully-modified estimates; constant and time trend included for all factor specifications

Table 5 shows results without the time-trend and can be similarly interpreted. Comparing with the previous table, the fully modified continuously updated estimator is larger in absolute value in most sectors but follows the previous estimates closely. There are only two abnormal sectors for which the full pass-through hypothesis cannot be rejected and for which the exchange rate effect contributes to very little of price variation. The

Philips and Moon estimator does a much poorer job than before, indicating the greater importance of cross sectional dependency when no time-trend is included. Finally, the model residuals appear to be much less stationary then the previous case, suggesting that spurious regression may now be a serious issue. Overall, there is enough evidence to conclude that the time-trend specification is the superior one. As suggested before, it is likely that other permanent shock disturb the long-run relationship of interest, and the deterministic trend proxies for them. From this point on, we consider only the coefficients estimated under the assumption of country specific constant and time-trend.

The inclusion of additional common factors preserves most of the results, apart from low technology sectors. Information criteria based on Bai and Ng (2002) provide no strong indication on the number of factors, with one or three factors being the preferred choice depending on the criteria; in any case, information criteria do not have robust properties in small finite samples. Table 6 reports the coefficient estimates and standard errors with as much as three common factors. Except for low technology ones where the single factor assumption seems to be essential, the pattern of results is fairly robust to the assumption. Without clear indication to the contrary by information criteria, parsimony and structural interpretation lead us to impose the single factor specification for all sectors.

5. Discussion

The main motivation for estimating long-run parameters was to improve on the previous panel literature, possibly obtaining a clearer picture on pricing-to-market behavior. The long-run coefficients are indeed much less dispersed then short-run analogues of the traditional panel literature reviewed before. Additionally, the sign pattern of negative mark-up effects significantly different from zero is more aligned with the aggregate evidence of incomplete pass-through which motivated the literature in the first place[5]. In this section, we show the estimated coefficients have interesting patterns and connections with microeconomic fundamentals.

[5] Aggregate time-series studies for the Brazilian exports points to incomplete small pass-through. Ferreira and Sanso (1999) found long-run pass-through ranging from 10% to 27%; Tejada and Silva (2005) typical estimates ranges from 14% to 34%. Our disaggregated estimates are somewhat higher, possibly suggesting aggregation bias in time-series studies.

The industrial classification scheme used in the paper is essentially based on the research and development activities in each industry which may have connections with preference and market structure parameters. High technology sectors are constantly developing new product varieties to attend very specific consumer needs - for example, consumers of medical appliances and pharmaceuticals very often have few substitution possibilities. As for the market, competition is not expected to be high if measured by the average mark-ups in the industry which reflects large market shares among key participants. On the opposite extreme, low technology industries represent consolidated business with a large number of players offering very substitutable commodities. There is indeed econometric evidence supporting this informal argument. Barroso (2009) found technological intensity is positively associated with product differentiation and supply elasticity.

Table 7. Long-run margin and pass-through effects by technological intensity

		Mark-up	Pass-through		
High	Pharmaceuticals	-0,40	0,60		
	Medical, precision and optical instruments	-0,26	0,74	0,672	
Medium-high	Eletrical machinery and apparatus	-0,23	0,77		
	Machinery and equipment	-0,38	0,62		
	Chemicals excluding pharmaceuticals	-0,46	0,54		
	Motor vehicles, trailers and semi-trailers	-0,53	0,47	0,597	
Medium-low	Rubber and plastics products	-0,46	0,54		
	Non-metallic mineral products	-0,40	0,60		
	Basic metals and fabricated metal products	-0,53	0,47	0,539	
Low	Food products, beverages and tobacco	-0,46	0,54		
	Wood, pulp, paper, paper products	-0,59	0,41		
	Other Manufacturing	-0,38	0,62		
	Textiles, leather and footwear	-0,43	0,57	0,536	
					0,58

Notes: (i)The pass-through equals one plus the mark-up reduction. (ii) Summary measures are means of sector estimates by technological intensity.

Table 7 calculates simple averages of pass-through coefficients for each level of technological intensity. At least for this sample, it appears that pass-through is increasing with technology. It is suitable to add some theoretical underpinning to these observations. For that matter, is not hard to argue that high substitution and low shares lead to lower pass-through. In Dornbush (1987), high substitution is complementary to competitors' responses to expected aggregate price changes in the industry, and each firm is forced to adjust markups more strongly. Atkeson and Burstein (2008) show that mark-ups are sensitive to market-shares, the more so for higher elasticities of substitution. Since mark-shares reflect differences in marginal costs the result follows. In both models, high market shares increases the influence on the sector price and therefore the pass-through that can be implemented. Of course, the actual details of the arguments leading to these interactions depend on the precise market structure and demand schedule assumed by the authors. But the pattern suggested by Table 7 and the underlying connection with microeconomic fundamentals argued for in the last paragraph both support models with this properties.

Another interesting pattern is the negative relationship between the sector share in total manufacture exports and the degree of pass-through. Indeed, sector share explains 30% of the variance of estimated coefficients. This might be indicative of pooling bias that need further investigation. Indeed, with fixed costs, domestic exporters would serve the most profitable destination markets first, where they can sustain a higher market share. But a sector with low participation in total exports is not highly developed and thus involves only the most profitable destinations. As a result, the pooled coefficient may give too much weight to low market share countries in traditional, low technology sectors. As a matter of fact, a similar argument could lead one to conclude that market share could be positive related to technological intensity. The possibility of pooling bias due to self-selection into destination countries should be further investigated. But care should be taken, because low market shares are also indicative of trade barriers which are of little consequence to the selection argument.

6. Conclusion

There is strong evidence of long-run pricing-to-market behavior by Brazilian exporters. Approximately 58% of an exchange rate appreciation would be passed-through to foreign consumer prices, with Brazilian exporters absorbing a 42% loss through reduced mark-ups. The degree of pass-through is positively related to the technological intensity of the industrial sector, a sensible pattern given the lower substitution between product varieties in high technology sectors. These results cannot be attributed to sticky-price constraints which by definition are not binding in the long-run. Therefore, the significant long-run effects support market structure explanations of incomplete pass-through and deviations from purchasing power parity, with possible normative consequences for trade and exchange rate policy.

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Gains from Imported Varieties in the Brazilian Economy

Abstract

This paper estimates welfare gains from new varieties of imported goods in Brazil since the 90s taking into full account preference and technology heterogeneity, in the sense of how easily substitutable are varieties of different goods to each other. Elasticities of substitution are estimated for a broad spectrum of consumer and intermediate goods using Feenstra (1994) method, which explores heteroscedasticity in the panel dimension to obtain identification of the structural parameters. An extension of Broda and Weistein (2006) aggregation model allowed productivity effects of intermediate goods to have an impact on welfare calculations. The welfare calculation points to significant gains from new varieties of imported goods, of the order of 0.5% of gross domestic product, most of which due to intermediate goods. Estimated elasticities of substitution are shown to have systematic relations to market structure indicators. They also point to significant aggregation bias in the literature's estimates of aggregate elasticities of substitution and aggregate price-indexes.

JEL Classification: C52, D60, F12, F40

Keywords: elasticity of substitution; gains from trade; import price index; imported variety; intermediate goods; heterogeneity.

1. Introduction

The greater availability of varieties is an important source of gains from trade, particularly when consumers or firms are less willing to substitute one variety for another [Krugman (1978), Ethier (1982)]. Indeed, elasticities of substitution between goods are crucial inputs for most positive and normative questions in open economy macroeconomics and international trade. They are usually estimated for broad categories, thus relying on strong homogeneity restrictions. Since the microeconomic evidence points to heterogeneity, some authors [Broda and Weinstein (2006), Imbs and Méjan (2008)] have been revisiting the aggregate exercises from more solid foundations. Building on ideas by Feenstra (1994), the new approach obtains sensible estimates of product level elasticities of substitution within a theoretical model that allows straightforward aggregation. This paper follows the approach to answer the question: what was the welfare impact from the changing varieties of imported goods since the opening up of the Brazilian economy in the 90s? Also of interest is if there are any discernible patterns in the elasticities of substitution estimates underlying the welfare impact calculations and what are their implications for aggregate import elasticities of substitution, relevant for counter-factual analysis.

Feenstra (1994) solved the problem of evaluating the welfare impact of a changing set of goods by, first, developing an exact index number theory and, second, advancing an econometric method to estimate the required preference or technology parameters. The economic argument is straightforward; since the reservation price for a constant elasticity of substitution specification is infinite, it is sensible to take to this limit, in the usual exact price index formula, the prices of unavailable goods. As shown in Feenstra (1994, 2006), this leads to an exact index, with the elasticity of substitution weighting the impact of new or disappearing varieties. In order to estimate the elasticity of substitution, the author confronts the classical endogeneity problem: prices and quantities are determined not only by consumption or firm input choices but also by supply decisions. In a moment of inspiration, Feenstra developed a panel framework in which it is possible to explore heteroscedastic demand and supply shocks along varieties to form consistent estimators for the structural parameters. Under the Armington assumption of goods differentiated by the country dimension, the author applied his method to a small sample of imported

commodities; the point was to evaluate if price effects in trade equations were underestimated due to unnoticed variation in the actual price index, which was the result for some of the goods analyzed by the author.

The method was extended to the many commodities case by Broda and Weinstein (2006), who were able to calculate an aggregate import price index with the appropriate correction for variety change. Note that absolutely new goods pose an obstacle to Feenstra's approach, which handles changing varieties only within goods categories with some trade data in both periods under consideration. Broda and Weinstein circumvent the issue by aggregating goods in broad categories whenever necessary. The authors also embed the import decision in the wider consumer decision problem, and are thus able to make aggregate welfare inferences. The convenient price index formula actually allowed the authors to calculate the compensating variation to improved variety, that is, a money metric equivalent of the welfare improvement obtained with the expanded set of goods. The gains from variety for the United States from 1972 to 2001 are estimated as 2.6 percent of 2001 income. An important shortfall in these calculations, acknowledged by the authors, is the bundling of commodities in a single consumer good category. This means gains from variety of imported intermediate and capital goods receive no separate treatment.

The productivity effect from new intermediate goods was the focus of a parallel growth literature. Indeed, Feenstra and Markusen (1992) anticipated most of the index number theory with expanding varieties in the context of firm decisions with the presence of differentiated intermediate goods; they also studied the consequences of variety expansion for growth in aggregate productivity. In a landmark paper, Romer (1993) was early to point out that fixed costs of introducing new intermediate goods from abroad implied in large welfare benefits from import tariff reductions, arguably much larger than the second order gains from lower distortions in existing markets which were the center of attention at the time. Taking these concerns into account, Klenow and Rodriguez-Clare (1997) calibrate a static general equilibrium model for Costa Rica where a 10% tariff reduction in intermediate and consumption goods may improve welfare in 2 percent, 1.5 percentage points of which result from expanded varieties – the welfare metric is a compensating variation relative to the base period income. Most of the variety gains in their counter-factual experiments stems from intermediate goods, for which the elasticity of

substitution, estimated with the usual time-series methods, was imputed to be relatively low. Rutherford and Tarr (2002) build a model with capital accumulation where the transition dynamics between steady-states after a 10% import tariff cut in intermediate goods leads, in the very long run, to a 10% welfare increase in present value terms, mainly due to the entry of new intermediate producers.

The results from this literature depend on estimates of aggregate elasticities of substitution between goods from home and foreign countries, the country of origin being the usual concept of variety. Most econometric methods indentify the parameters of interest from the responsiveness of aggregate quantities to changes in aggregate prices, which usually results in very low elasticities of substitution relative to microeconomic studies. Imbs and Méjan (2008) document aggregation bias originated from the positive correlation between elasticities of substitution and shares in the enclosing category. They argue that aggregate elasticities of substitution relevant for representative agent models are actually much larger than the usual estimates[6]. Thus, the gains from new imported varieties derived from general equilibrium models with endogenous introduction of new goods may be lower than previously thought.

This paper makes several contributions to the literature. An important advance is to extend Broda and Weinstein (2006) method to account for the productivity effect of importing new kinds of intermediate goods. The extension is particularly relevant for a developing country such as Brazil in the nineties. The paper advocates a chained index method which reduces aggregation due to the absolutely new good problem mentioned above and allows more accurate figures on year-to-year gains from trade. On the econometric side, an improvement in robustness is obtained from coupling Feenstra's estimator and a parametric bootstrap procedure. Given the large number of models and the likely presence of undetected outliers, it is important to have a robust procedure. The estimated elasticities of substitution are confronted with known patterns documented by Broda and Weinstein (2006); also, an attempt is made to uncover new regularities. The paper clarifies the relation between the welfare gain specification and Imbs and Méjan

[6] As stressed by the authors, the muted response of aggregate quantities to prices stems from relatively large price variation coming from low elasticity goods, and not a low representative substitution parameter.

(2008) specification, thus allowing for consistent aggregation of estimated elasticities of substitution.

There are a number of substantive contributions, more closely concerned with Brazil, with potential application in open economy macroeconomic research. These range from aggregate import price index corrected for the introduction of new goods, to aggregate elasticities of substitution for imports free from aggregation bias. The elasticities estimates are particularly relevant for calibrating general equilibrium models of the Brazilian economy. As an appraisal of this estimates, the paper contrasts them with the available state of the art time-series estimates of Armington elasticities. In addition, disaggregated elasticities parameters themselves are important inputs for many kinds of frontier research - for example, Chaney (2008) notes that the elasticity of substitution provides conflicting incentives for the extensive and intensive margin of bilateral trade flows, and used Broda and Weinstein (2006) estimates to explore the issue empirically. The main substantive contribution, though, is a money metric measure of the welfare gains from new varieties of imported goods for the Brazilian Economy in the 1989 to 2008 period, which has not been addressed by previous research.

2. Economy

2.1. Preferences and Technology

Each consumer good is available in many varieties, which represent the country where the good was produced. As in Broda and Weinstein (2006), consumer preferences are assumed to be separable in imported and domestic varieties of every good. This affords great simplification, since the impact of sourcing a good from a new country does not depend on domestic production of the good at all, but only on alternative foreign suppliers. Further implications for aggregation and welfare impact inference will be discussed bellow. We also assume separable preferences between sets of varieties, and adopt a nested constant elasticity of substitution framework.

Imported consumer good k may be available in varieties $i \in I$, which refer to foreign countries. Preferences over the consumption of varieties $(m_{kit})_{i \in I}$ of imported good k at time period t are represented by the function

$$(1) \qquad m_{kt} = \left(\sum_{i \in I} b_{kit}^{\frac{1}{\sigma_k}} m_{kit}^{\frac{\sigma_k - 1}{\sigma_k}} \right)^{\frac{\sigma_k}{\sigma_k - 1}} \quad , \sigma_k > 1$$

where σ_k is the constant elasticity of substitution for varieties of good k and b_{kit} is the time varying quality parameter for variety i of good k. Moving up, preferences over the consumption of sets of varieties of imported goods are represented, in terms of consumption aggregates $(m_{kt})_{k \in K}$, by the function

$$(2) \qquad m_t = \left(\sum_{k \in K} m_{kt}^{\frac{\gamma - 1}{\gamma}} \right)^{\frac{\gamma}{\gamma - 1}} \quad , \gamma > 1$$

where γ is the elasticity of substitution for imported goods. Finally, preferences over all goods and varieties are defined in terms of the import aggregator m_t and an arbitrary domestic goods aggregator d_t,

$$(3) \qquad u_t = \left(d_t^{\frac{\kappa - 1}{\kappa}} + m_t^{\frac{\kappa - 1}{\kappa}} \right)^{\frac{\kappa}{\kappa - 1}} \quad , \kappa > 1$$

where κ is the constant elasticity of substitution between domestic and imported goods.

Intermediate goods are also available in many varieties, according to the country of origin. Feenstra and Markusen (1992) deduce the properties of the aggregate gross production function from the decentralized optimization problems of many possibly non-competitive sectors; in particular, they study the impact of new intermediate goods varieties. The present specification is an example of this general function. As in the consumer side of the model, imported intermediate goods are supposed to enter as

separable inputs in the economy gross production function and the set of varieties for each good are also separable from each other.

Each imported intermediate good j may be available in varieties $i \in I$. The economy bundles imported intermediate varieties $(x_{jit})_{i \in I}$ in period t to produce the aggregate intermediate good x_{jt} using the production function

(4) $$x_{jt} = \left(\sum_{i \in I} a_{jit}^{\frac{1}{\sigma_j}} x_{jit}^{\frac{\sigma_j - 1}{\sigma_j}} \right)^{\frac{\sigma_j}{\sigma_j - 1}} \quad , \sigma_j > 1$$

where σ_j is the constant elasticity of substitution for varieties of good j and a_{jit} is the time varying quality parameter for variety i of good j. The intermediate goods $(x_{jt})_{j \in J}$ are further aggregated into a single imported intermediate good x_t with production function

(5) $$x_t = \left(\sum_{j \in K} x_{jt}^{\frac{\delta - 1}{\delta}} \right)^{\frac{\delta}{\delta - 1}} \quad , \delta > 1$$

where δ is the constant elasticity of substitution for intermediate goods. The economy takes as input this imported intermediate good x_t, an aggregate of domestic intermediate goods z_t and other primary factors l_t to obtain gross income with the production function

(6) $$y_t = l_t^{1-\alpha} \left(z_t^{1-\omega} x_t^{\omega} \right)^{\alpha} \quad , \omega, \alpha \in (0,1)$$

Home intermediate goods are produced from primary factors with constant returns to scale.

2.2. Equivalent Variation

In this section, after restating the index number results from Feenstra (1994), we deduce the equivalent variation due to variety change taking into account both consumer and intermediate goods. This amounts to a simple extension of Broda and Weinstein (2006) to accommodate the production technology introduced above.

The representative consumer faces prices $p_t=(p_{kt})_{k\in K}=(\{p_{kit}:i\in I_{kt}\})_{k\in K}$, where $I_{kt}\subset I$ is the subset of countries from which good k may be imported in period t. The set of available varieties in each period is taken as given. From equation (1), the unit cost function associated with consumer good k is

$$c_m(p_{kt},b_{kt},I_{kt})=\left(\sum_{i\in I}b_{kit}{p_{kit}}^{1-\sigma_k}\right)^{\frac{1}{1-\sigma_k}} \tag{7}$$

The exact price index for good k given prices and consumer choices, supposing constant quality parameters b and a fixed set of available varieties $\check{I}_k$, is the well known Sato-Vartia exact price index,

$$p_m\left(p_{kt},p_{kt-1},m_{kt},m_{kt-1},\check{I}_k\right)=\frac{c_m\left(p_{kt},b,\check{I}_k\right)}{c_m\left(p_{kt-1},b,\check{I}_k\right)}=\prod_{i\in\check{I}_k}\left(\frac{p_{kit}}{p_{kit-1}}\right)^{w_{kit}} \tag{8}$$

where w_{kit} is the normalized log-change share in the good's expenditure defined by

$$w_{kit}=\frac{s_{kit}-s_{kit-1}}{lns_{kit}-lns_{kit-1}}\Bigg/\sum_{i\in\bar{I}_k}\frac{s_{kit}-s_{kit-1}}{lns_{kit}-lns_{kit-1}},\qquad s_{kit}=\frac{p_{kit}m_{kit}}{\sum_{i\in\bar{I}_k}p_{kit}m_{kit}} \tag{9}$$

Let I_k be a subset of varieties from good k available in period t and t-1, for which quality parameters have remained constant. We may refer to this subset as the common varieties. As shown by Feenstra (1994), the exact price index for gook k is given by

$$\begin{aligned}&\pi_m(p_{kt},p_{kt-1},m_{kt},m_{kt-1},I_{kt},I_{kt-1})\\&\qquad=P_m(p_{kt},p_{kt-1},m_{kt},m_{kt-1},I_k)\left(\frac{\lambda_{kt}}{\lambda_{kt-1}}\right)^{\frac{1}{\sigma_k-1}}\end{aligned} \tag{10a}$$

where $\lambda_{k\tau}$ is the share of common varieties for $\tau\in\{t,t\text{-}1\}$, that is

(10b) $$\lambda_{k\tau} = \frac{\sum_{i \in I_k} p_{kit} m_{kit}}{\sum_{i \in I_{k\tau}} p_{kit} m_{kit}}.$$

The intuition for the result is taking the prices of unavailable varieties to infinity in the ratio of unit costs with different sets of varieties; the decomposition in equation (10) then follows naturally. As is evident in the formula, new varieties reduce the lambda ratio and therefore the exact price index; small elasticities of substitution and large shares of these varieties in total consumption both contribute to a larger reduction.

Aggregation is straightforward. Indeed, as in Broda and Weinstein (2006), suppose there is a collection $I_K=(I_k)_{k\in K}$ of subsets of common varieties for each good in a constant set of goods K. Apply the Sato-Vartia exact price index to the aggregate consumption of imported goods m_t shown in equation (2). Now, the unit cost function is the correct concept of aggregate import price; and we obtain the exact price index number for imported goods:

(11a) $$\pi_M(p_t, p_{t-1}, m_t, m_{t-1}, I_{Kt}, I_{Kt-1}) = \prod_{k\in K} P_m(.\,, I_k)^{w_{kt}} \prod_{k\in K} \left(\frac{\lambda_{kt}}{\lambda_{kt-1}}\right)^{\frac{w_{kt}}{\sigma_k - 1}}$$

where w_{kt} is the normalized log-change share in import expenditure, and the arguments in the price index formula for each good where suppressed. At times, it will be convenient to use the more compact notation

(11b) $$\pi_M = P_M(I_K)\Lambda_M(\sigma_K)$$

By the same argument, the aggregate consumer price index associated with equation (3) is

(12) $$\pi_U = \frac{c_{Ut}}{c_{Ut-1}} = \left(\frac{p_t^d}{p_{t-1}^d}\right)^{w_z} P_M(I_K)^{w_m} \Lambda_M(\sigma_K)^{w_m}$$

where w_d and w_m are log-change share of domestic and imported goods in consumption expenditure, and the first equality is the ratio of consumption unit costs. For emphasis, we summarize the results in a proposition.

Proposition 1 [Feenstra (1994), Broda and Weinstein (2006)]. Suppose there is a common set of varieties I_k. Then the exact price index for good k is given by (10). If in addition there is a collection of common varieties $I_K=(I_k)_{k\in K}$, then the exact aggregate price index is given by (11).

The firm's cost minimization problem for the production of the composite intermediate good is entirely analogous to the consumer problem. From Propostion 1, with the evident notation, the exact price index for imported intermediate goods is given by

(13) $$\pi_X\left(p_t^x, p_{t-1}^x, x_t, x_{t-1}, I_{Jt}, I_{Jt-1}\right) = \prod_{j\in J} P_x\left(.,I_j\right)^{w_{jt}} \prod_{j\in J}\left(\frac{\lambda_{jt}}{\lambda_{jt-1}}\right)^{\frac{w_{jt}}{\sigma_j-1}}$$

To obtain the aggregate price index for the economy´s gross output, apply the usual Cobb-Douglas index,

(14) $$\pi_Y = \frac{c_{Yt}}{c_{Yt-1}} = \left(\frac{p_t^l}{p_{t-1}^l}\right)^{1-\alpha}\left[\left(\frac{p_t^z}{p_{t-1}^z}\right)^{1-\omega} P_X\left(I_J\right)^{\omega} \Lambda_X\left(\sigma_J\right)^{\omega}\right]^{\alpha}$$

Now, total production costs are equal to unit costs times production. Therefore,

(15) $$\frac{y_t}{y_{t-1}} = \frac{C_{Yt}}{C_{Yt-1}}\Big/\frac{c_{Yt}}{c_{Yt-1}} = \frac{C_{Yt}}{C_{Yt-1}}(\pi_Y)^{-1}$$

We have all the ingredients to compute the equivalent variation. Indeed given consumer income in period t is $I_t = u_t C_U(p_t) = (1-\alpha)y_t + z_t$, the equivalent variation relative to the initial period income is

(16) $$\frac{EV}{I_0} + 1 = \frac{u_1 C_U(p_0)}{u_0 C_U(p_0)} = \frac{I_1}{I_0}\frac{C_U(p_0)}{C_U(p_1)} = \frac{y_1}{y_0}\frac{C_U(p_0)}{C_U(p_1)} = \frac{C_{Yt}}{C_{Yt-1}}(\pi_Y\pi_U)^{-1}$$

The third equality used the production functions defined above so that z may be eliminated from the income ratio. Just divide the numerator and denominator by $(1-\alpha)y_1$ and use Cobb-Douglas for final goods, linearity on labor for intermediate goods, as labor as numeraire. To emphasize changes available varieties, rewrite the expression as

$$\frac{EV}{I_0}+1=EV\left(I_K,I_J\right)\left[\Lambda_X\left(\sigma_J\right)^{\alpha\omega}\Lambda_M(\sigma_K)^{w_m}\right]^{-1} \tag{17}$$

where $\mathrm{EV}\left(\mathrm{I_K},\mathrm{I_J}\right)$ depends only on the collection of common varieties and the last term is a measure of the welfare impact of new varieties.

Proposition 2. The equivalent variation may be decomposed in two terms, one of them measuring the welfare impact of variety change, as in equation (17).

The welfare measure just derived depends on the elasticity of substitution for many disaggregated goods. To estimate these parameters, the paper applies Feenstra's (1994) method, briefly reviewed in next section. We take some time to show the method is robust to tariffs and exchange rate pass-through decisions by firms.

3. Identification

From the cost minimization problems faced by consumers and firms, demand at period *t* for the country/variety *i* of a particular good *k* may be written in terms of their share s_{kit} in expenditure,

$$s_{kit}=\left(\frac{p_{kit}}{c_{kt}}\right)^{1-\sigma_k}(\beta_{kit})^{\sigma_k-1} \tag{18}$$

Consistent with our dataset, suppose observed prices $\tilde{p}_{kit}$ are measured in dollars and net of tariffs. Their relation with consumer prices is

(19) $$\tilde{p}_{kit} = \frac{f_t p_{kit}}{\tau_{kt}} = \frac{(d_{it}/e_{it})p_{kit}}{\tau_{kt}}$$

where τ_{kit} is the discount factor for tariffs and f_t is the exchange rate in dollar per unit of the importer's currency. As a matter of notation, d_{it} and e_{it} are the foreign country *i* exchange rates with respect to dollar and importer country currency. We also define the observed expenditure shares as

(20) $$\tilde{s}_{kit} = \frac{\tilde{p}_{kit}x_{kit}}{\langle \tilde{p}_{kt}, x_{kt}\rangle} = \frac{f_t}{\tau_{kit}} \frac{p_{kit}x_{kit}}{\langle p_{kt}, x_{kt}\rangle} \frac{\langle p_{kt}, x_{kt}\rangle}{\langle \tilde{p}_{kt}, x_{kt}\rangle} := \frac{f_t}{\tau_{kt}} s_{kit}\mu_{kt}$$

Substituting in (18) and taking logarithms

(21) $$ln\tilde{s}_{kit} = \Phi_{\mathrm{kt}} + (1-\sigma_k)ln\tilde{P}_{kit} + \underbrace{(\sigma_k - 1)\beta_{kit}}_{\varepsilon_{kit}}$$

with $\Phi_{\mathrm{kt}} = (\sigma_k - 1)lnc_{kt} + ln\mu_{kt} + \sigma_k lnf_t - \sigma_k ln\tau_{kt}$. Now apply first-differences with respect to time, followed by differencing with respect to a reference country i_k^*, to eliminate Φ_{kt}. Denote this double differencing by Δ^*, and we obtain the final expression for demand

(22) $$\Delta^* ln\tilde{s}_{kit} + (\sigma_k - 1)\Delta^* ln\tilde{p}_{kit} = \Delta^*\varepsilon_{kit}$$

Note that eliminating tariffs from the demand equation required an equal treatment assumption on foreign countries. If tariffs or any other transaction costs vary with the country of origin, they will be encapsulated in the error ε_{kit}.

Firms' pricing decision p_{ki}^*, tariffs and the relevant exchange rate determine prices from the supply side

(23) $$p_{kit} = \tau_{kt}e_{it}p_{kit}^*$$

Other than that, the economy has no assumption on the supply side pricing decision. For the purpose of econometric identification, consider the following reduced form decision rule

$$lnp_{ki}^{*} = \omega_k^{\tau} ln\tau_{kt} + \omega_k^{e} lne_{it} + \omega_k lnx_{kit} + v_{ki} \tag{24}$$

where ω_k is the inverse supply elasticity and the first two coefficients determine the pass-through from tariffs and exchange rates, the simplest case being full pass-through (when $\omega_k^{\tau} = \omega_k^{e} = 0$). Substituting back in (23), after using (19) and (20), we get

$$ln\tilde{p}_{kit} = \Psi_{\mathrm{kt}} + \omega_k ln\tilde{s}_{kit} - \omega_k ln\tilde{p}_{kit} + \underbrace{v_{ki} + (1 + \omega^{E}) lne_{it}}_{\delta_{kit}} \tag{25}$$

where $\Psi_{\mathrm{kt}} = \omega^{\tau} ln\tau_{kt} + \omega ln\langle \tilde{P}_{kt}, C_{kt}\rangle + f_t$. Application of Δ^* leads to the final expression for the supply side

$$(1 + \omega_k)\Delta^* ln\tilde{p}_{kit} - \omega_k \Delta^* ln\tilde{s}_{kit} = \Delta^* \delta_{kit} \tag{26}$$

As shown in Feenstra's (1994), multiplying (22) and (26) and solving for the price variable results in a regression equation that may be consistently estimated with deterministic instruments (esp. country of origin dummies). The instrumental variables estimator is equivalent to a weighted regression in the time averages of each country/variety included in the regression, with the number of periods as weights. As this equivalent representation makes clear, the asymptotic exercise considered by Feenstra is conducted in the time dimension. A constant was also included in the regression to capture measurement error in the price variable.

The first crucial identifying assumption is independence between δ and ε at all time periods and varieties. Equations (22) and (26) make clear this either requires homogeneous tariffs across varieties of the same product or full tariff pass-through to the importing country. In particular, the exchange-rate pass-through behavior of firms is irrelevant. The standard heterogeneous firm model under constant elasticity of substitution preferences implies full tariff and exchange rate pass-through; therefore, it is a convenient model to have in mind for identification purposes.

The second and last crucial assumption is heteroscedastic errors across varieties, which will guarantee the rank condition for the instrumental variables estimator or, in other words, will avoid colinearity in the equivalent weighted regression. In any case, there is a bijective mapping between regression coefficients and the parameters in equation (22) and (26), guaranteeing identification of the relevant parameters[7].

4. Data

The data is organized in panels, one for each imported product. Products are defined at six digits in the Harmonized System of commodity classification. Standard international conversion tables were adopted in years where the harmonized system suffered revision. Occasionally, resolution rules where appended to the conversion tables, always observing the principle of favoring the most traded item in the most recent period. The adoption of conversion tables increases sample size in the time dimension and reduces the necessary aggregation in welfare calculations due to spurious new products; it appears to be a significant methodological improvement relative to the previous literature. The result is a data set with almost 4,000 panels and welfare calculations based on very much disaggregated data. It is worth noticing that the panel structure of the dataset is not explored for identification purposes beyond making more plausible the rank condition mentioned in the previous section.

Considering the sample size, each panel extends for at most twenty years (1989-2008) and at most ninety countries of origin, where each country represents a different variety. The time dimension is very small for the asymptotic argument necessary for identification. For this reason, point estimates are not to be taken as seriously as the aggregates built from them. We will report and be concerned with these aggregates when considering patterns and economy wide effects. In any case, confidence intervals will always be calculated with bootstrap procedures meant to approximate finite sample properties; never with asymptotic limits.

[7] An alternative identification strategy is to fix the time dimension and consider averages across at least two heteroscedastic groups of countries, with asymptotics now running in the number of countries (admitting some strictly positive probability limit for the shares of countries in each group). We leave the suggestion for further study.

The panels are unbalanced due to the changing varieties in the import bundle. Figures 1 and 2 summarize the information on variety in our data set. The first one shows the evolution of the mean number of countries/variety per product. The second one presents the normalized number of products by country, where the mean count across countries was taken as the base for the normalization. Both figures show significant entry and exit of varieties in the Brazilian import bundle. In the beginning of the nineties there was strong import diversification, followed by a reduction period starting at 1998, and a latter recovery from 2003 on. Of course, the welfare implications of these strong trend and cycles in import varieties are the main concern for this paper.

Figure1 – Mean country oforigin per imported products, annual data, 1989-2008
Product = 6 digitsHarmonized System

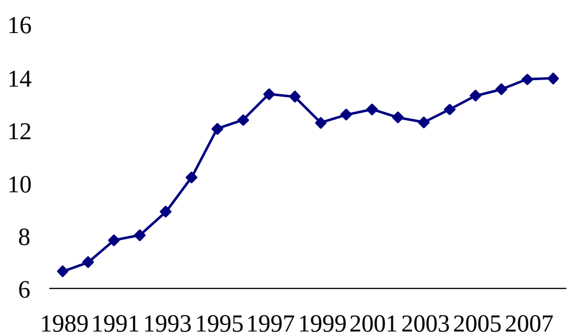
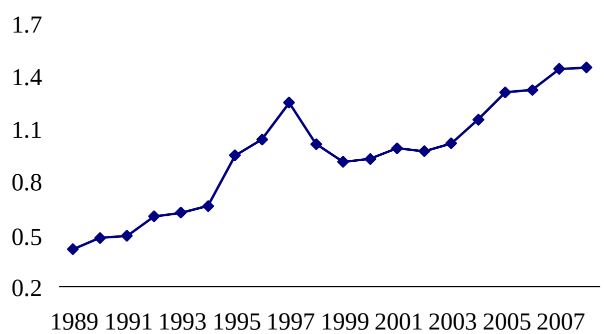


Figure 2 – Mean number of products imported per country of origin, annual normalized data, 1989-2008
Product = 6 digits Harmonized System



5. Results

5.1. Elasticities of Substitution

We apply several commodity classification schemes to uncover patterns related to market structure, technological intensity, economic sector and economic category. In general, there are discernible regularities in the elasticities of substitution estimates which both corroborate and extend previous results in the literature.

Considering market structure, we followed Broda and Weinstein (2006) and applied Rauch's (1999) classification of commodities in the following categories (in order of increasing degree of price and product homogeneity): differentiated, reference priced or exchange traded commodities. Results are summarized in Table 1 by the median and the

trimmed mean estimates for each class of good. Bootstrapped 90% confidence intervals are shown in parenthesis. All the bootstrap exercises conducted were parametric, admitting a heteroscedastic normal distribution for the errors.

The ordering is natural. Exchange traded goods are characterized by varieties easy to substitute one for another. Goods with prices negotiated with a public reference price are less substitutable. Finally, differentiated goods, the ones priced with most discretion by suppliers, exhibit low elasticity of substitution. The differences are significant, with p-values at zero in pair wise bootstrap permutation tests. Moreover, it is consistent with results from Broda and Weinstein obtained with a completely different data-set.

As also shown in Table 1, industry supply is more elastic the more homogeneous the supplied goods are, again in line with prior expectations. These results serve as important first checks on the plausibility of the elasticity estimates. They also favor the view of a positive association between market structure as measured by the supply elasticity and consumer substitution behavior.

Table 1. Elasticities of Substitution and Inverse Supply by Market Structure

	Substitution (σ)		Inverse Supply (ω)		
Market structure [/1]	Median	Mean [/2]	Median	Mean [/2]	#Products
(i) Differentiated	**3.99**	**5.55**	**0.48**	**0.70**	2337
	(3.91, 4.07)	(5.41, 5.73)	(0.46, 0.49)	(0.66, 0.74)	
(ii) Reference priced	**4.06**	**6.50**	**0.31**	**0.52**	1048
	(3.96, 4.16)	(6.27, 6.85)	(0.30, 0.34)	(0.47, 0.56)	
(iii) Exchange Traded	**4.74**	**8.43**	**0.22**	**0.33**	276
	(4.49, 4.98)	(7.77, 9.64)	(0.19, 0.25)	(0.25, 0.41)	

[/1] Uses Rauch's (1999) commodity classification by market structure

[/2] Trims 5% percentile of extreme elasticities from both tails in all bootstrap samples

To study the relationship between technological intensity and the parameters of interest, the paper applies a classification system developed by OCDE and often used by official institutions. It is primarily based on the intensity of research and development expenditure on a sector, with commodities assigned to sectors. Results are summarized in Table 2 by the median and the trimmed mean estimates for commodities classified from high technology sectors to low technology sectors. As before, bootstrapped 90% confidence

intervals are shown in parenthesis. The intervals also serve as summary information on the results of bootstrap permutation tests for differences between averages in pairs of categories. In general, null intersection in the confidence intervals leads to zero p-values for the null of equal averages, while even the smaller intersection in the table leads to p-values higher than .10, therefore supporting the null of equal averages.

With this in mind, low technology sectors are clearly associated with higher elasticities of substitution then other sectors. But further separation of sectors as medium-low, medium-high or high technology leads to no significant distinction in elasticities of substitution. As for supply elasticities, they are decreasing in technological intensity, although with no definite ordering between the medium technology categories. This result also favors a positive association between competition, as measured by supply elasticity, and the degree of substitution in consumption. These regularities related to technological intensity are new to the literature, although certainly in line with prior expectations.

Table 2. Elasticities of Substitution and Inverse Supply by Technological Intensity

	Substitution (σ)		Inverse Supply (ω)		
Tech Intensity [/1]	Median	Mean [/2]	Median	Mean [/2]	#Products
(i) High	**3.83**	**5.21**	**0.60**	**0.83**	387
	(3.67, 4.02)	(4.93, 5.55)	(0.55, 0.67)	(0.67, 1.00)	
(ii) Medium-High	**3.82**	**5.45**	**0.43**	**0.66**	1290
	(3.74, 3.89)	(5.27, 5.69)	(0.41, 0.45)	(0.61, 0.70)	
(iii) Medium-Low	**3.70**	**5.38**	**0.41**	**0.63**	735
	(3.60, 3.82)	(5.10, 5.64)	(0.39, 0.44)	(0.57, 0.70)	
(iv) Low	**4.67**	**7.19**	**0.35**	**0.53**	1085
	(4.53, 4.80)	(6.85, 7.51)	(0.33, 0.37)	(0.48, 0.58)	

[/1] OCDE's (1999) commodity classification by technology level in its industry

[/2] Trims 5% percentile of extreme elasticities from both tails in all bootstrap samples

With respect to economic sector, as argued by Imbs and Méjan (2008), the weighted average of substitution elasticities in a sector is a first order approximation to the aggregate substitution elasticity relevant for macroeconomic counterfactual exercises and modeling. The correct theoretical weight is shown to be the product's share in the enclosing sector. The authors documented a strong downward bias in aggregate estimates obtained from quantity index responses to the sector's price index, which they attribute to a

Table 3. Elasticity of Substitutuion by Industry: panel and reference estimates.

Industry	Panel /1	Reference /2	Corr	#	Industry (continuing)	Panel /1	Reference /2	Corr	#
Agriculture	**10.88**	**3.8**	0.10	74	Petroleum Refining and Petrochemical Ma	**3.23**	**0.18**	-0.2	14
	(6.7, 31.3)	(1.5, 6.0)				(2.8, 5.0)	(-0.1, 0.5)		
Coal Mining	**8.66**	**-2.38**		2	Chemical Products Manufacturing	**6.07**	**1.51**	0.02	634
	(4.4, 57.2)	(-3.7, -1.0)				(5.6, 8.0)	(0.3, 2.7)		
Petroleum and Gas Extraction	**13.81**	**0.6**		1	Pharmaceutical Manufacturing	**5.05**	**0.58**	-0.0	71
	(9.9, 25.5)	(-0.0, 1.2)				(4.1, 6.4)	(-0.0, 1.2)		
Nonmetallic Minerals Mining	**2.96**	**-2.38**		9	Plastics and Rubber Products Manufactur	**5.94**	**1.22**	-0.0	120
	(2.5, 5.1)	(-3.7, -1.0)				(4.7, 7.7)	(0.8, 1.5)		
Metallic Minerals Mining	**11.42**	**1.24**	**0.40**	32	Metals Production and Processing	**4.53**	**3.06**	-0.0	122
	(5.7, 28.7)	(0.6, 1.8)				(4.0, 5.0)	(2.5, 3.6)		
Food Manufacturing	**8.74**	**0.95**	-0.0	218	Metal Products Manufacturing	**6.66**	**0.47**	0.05	253
	(7.7, 10.2)	(0.7, 1.1)				(5.4, 7.7)	(-0.2, 1.2)		
Tobacco Processing	**9.63**	**2.47**	0.07	17	Electrical Equipment and Components Ma	**3.86**	**0.2**	-0.0	196
	(7.4, 14.3)	(1.0, 3.9)				(3.5, 4.3)	(-0.0, 0.4)		
Textiles Manufacturing	**10.12**	**2.34**		6	Machinery and Equipment Manufacturing	**4.17**	**1.84**	-0.0	244
	(5.9, 19.4)	(1.9, 2.7)				(3.9, 4.5)	(1.4, 2.2)		
Apparel and Accessories Manufacturing	**8.52**	**2.2**	**0.03**	289	Automobile, Truck and Bus Manufacturing	**4.07**	**5.28**	-0.1	146
	(6.9, 10.6)	(1.8, 2.5)				(3.7, 4.7)	(3.0, 7.5)		
Footwear and Leather Products Manufactu	**7.05**	**0.15**	0.03	188	Other Transportation Equipment Manuf.	**4.73**	**0.19**	-0.0	388
	(6.0, 8.6)	(-0.6, 0.9)				(4.4, 5.2)	(-0.0, 0.4)		
Wood Products Manufacturing	**6.66**	**1.58**	0.09	56	Furniture Manufacturing	**8.66**	**1.58**	0.13	56
	(5.6, 8.4)	(0.9, 2.1)				(5.7, 15.3)	(0.9, 2.1)		
Pulp and Paper Production	**10.13**	**0.51**	0.12	38	Miscellaneous Other Products Manuf.	**5.49**	**2.65**	0.02	44
	(5.8, 13.5)	(0.2, 0.8)				(4.0, 8.7)	(-1.1, 6.4)		

/1 Product level estimates obtained by Feenstra (1994) panel method; industry level aggregation is a weighted average as in Imbs and Mejan (2008). Extreme estimates were excluded before taking industry average in all bootstrap samples.

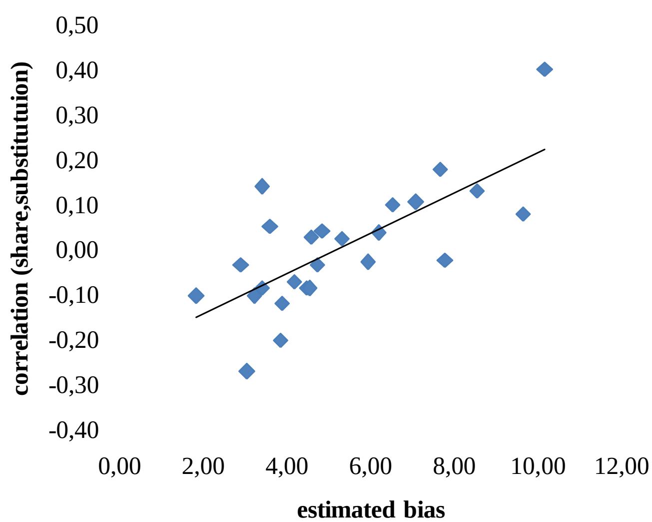
/2 Kume, Pedroso e Tourinho (2003) time-series method for industry leveldata. Considering the updated industry classification squeme used in this paper, the pairing exhibited in the table is only approximate.

disproportional weight on low elasticity components of the sector. In particular, constraining elasticities to be the same will build a downward bias whenever the most representative products from a sector turn out to have high elasticities of substitution. From another perspective, most of the aggregate price change will refer to small elasticity of substitution goods, which require more price adjustment to obtain market equilibrium, biasing time-series estimates.

Following Imbs and Méjan (2008), estimates for aggregate elasticity of substitution were obtained for a comprehensive list of Brazilian industrial sectors using the weighted average of the elasticities of substitution of products included in the sector. Table 3 reports the results with bootstrapped 90% confidence intervals bellow the estimates. For comparison, state of the art elasticity of substitution estimates obtained through traditional aggregated time-series exercises are shown in the same table - the correspondence being only approximated due to classification differences between sectors. As expected from previous studies, the microeconomic estimates taking heterogeneity into account are generally higher than macroeconomic estimates which implicitly impose homogeneity restrictions.

The correlation between elasticities of substitution and the relative share in the sector is also reported in the table (in the column named "corr"). The correlation is significantly different from zero only for Metallic Minerals and Mining sector and for Apparel and Accessories Manufacturing sector, and significantly positive in both cases. Despite this, there is an overall positive relation, illustrated in Figure 3, between point correlation estimates and the estimated aggregation bias, where the bias is defined as the difference between our estimate and the

Figure 3. Relation between aggregate elasticity of substitution estimated bias and intra-industry distribution of elasticities of substitution



reference figure. This relation lends support to Imbs and Méjan (2008) theoretical rationale for the aggregation bias. It also suggest that aggregate models using the hypothesis of a behavioral representative agent should probably revise their elasticity of substitution estimates upwards; and considerably so in some cases.

Our final result regarding patterns on elasticities of substitution refers to the economic category of the goods by main economic use. Consistent with our theoretical model used bellow to calculate welfare gains, goods were classified as either intermediate or consumer goods. Taking as starting point the international standard classification in Broad Economic Categories, capital goods were aggregated into intermediate goods and other durable commodities were aggregate into consumer goods. We take the previous results concerning sectors as positive evidence to estimating behavioral (that is, representative agent) aggregate elasticity of substitution as the weighted average of the aggregated products. Table 4 shows the trimmed mean, as a summary measure, and the weighted mean, as an estimator for the aggregate parameter of each category. As before, bootstrapped confidence intervals are shown bellow in parenthesis. The results (p-values) of permutation tests for the equality of the mean estimates for intermediate and consumer goods are also reported in the table. Intermediate goods have on average lower elasticities of substitution than consumer goods; however, for the purpose of macroeconomic counterfactual exercises, there is weak evidence for lower intermediate elasticities.

Table 4. Elasticities of Substitution by Economic Category (main end use)

Category		Mean [/1]	Weighted mean [/2]	#Products
(i) Intermediate goods		**5.67**	**6.46**	2786
		(5.55, 5.88)	(5.63, 7.50)	
(ii) Consumer goods		**6.98**	**8.30**	876
		(6.67, 7.37)	(6.43, 11.77)	
Test for equality		Mean [/1]	Weighted mean [/2]	
i=ii	(i<ii)	0.00	0.27	

Note: Durables and capital goods included in consumer and intermediate aggregates

[/1] Trims 5% percentile of extreme elasticities from both tails in all bootstrap samples

[/2] Estimate of aggregate import elasticity; uses only trimmed values for the average.

5.2. Gains from Variety

The crucial step for calculating gains from variety is to obtain the lambda ratios defined in equation (10). Note that each product should have a common set of varieties for both periods under consideration. This is an unfeasible requirement for a product not imported from any country before a certain date. Other situations may prompt the same problem, for example, radical changes in the sourcing countries of a good or changes in the six digits code system used as product key. The former problem was circumvented by the systematic application of standard international translation tables for every years of occurrence of a change in the six digits Harmonized System of commodity classification. But the definitive solution, as in Broda and Weinstein (2006), is to aggregate products by dropping two of more digits whenever we face the unfeasibility problem[8]. The elasticity of substitution estimate for the aggregate products is defined to be the weighted average of the available estimates of elasticities of substitution of its components.

We applied this procedure for each year in our sample, taking the immediately following year as the second period in the exact price index formula. The formula for the price index or, for that matter, the formula for the equivalent variation in Proposition 2 could in principle be chained to obtain the exact formula for further apart periods. However, the product aggregation scheme would have to be the same through all the chain. Since new products are more likely to appear the further apart are two periods, the aggregation scheme would tend to become too coarse, with heterogeneity arbitrarily reduced. To avoid this problem, we applied chained indexes without imposing aggregation consistency, so maximum heterogeneity is considered in each step of the calculation. An additional benefit from this strategy is a accurate figure of the dynamics of the gains from variety from year to year.

To gauge the relative impact of varieties on consumer and intermediate goods, it is useful to calculate the lambda ratio component of the chained price index for both economic categories. Figures 4 and 5 show the ratios, along with 90% bootstrapped

[8] The exact algorithm is the following: start with full digits; {if the product with the current number of digits is isolated, apply for a reduction in digits; if the product is not isolated and there is no applicant for reduction with the same initial digits as this product, define a product with the current digits}; continue the procedure, now with less digits.

confidence intervals. From 1989 to 2008, the cumulative upward bias in the associated price indexes is almost 5% for both categories. As originally pointed out by Feenstra (1994), disregarding this downward trend in import price indexes actually faced by consumers and firms may lead researches to overstate the effect of income variables on trade fluctuation. A surprising result illustrated in the figures is the contrasting behavior of consumer and intermediate goods in the beginning of the nineties. At the same time variety of consumer goods rapidly increased, intermediate goods varieties were becoming less widely available. These years are depicted in the literature as a period of intense trade liberalization, an assertion that must be qualified when it comes to the likely negative effects of the period's policies on intermediate goods varieties.

Figure 4. Consumption goods chained lambda ratio

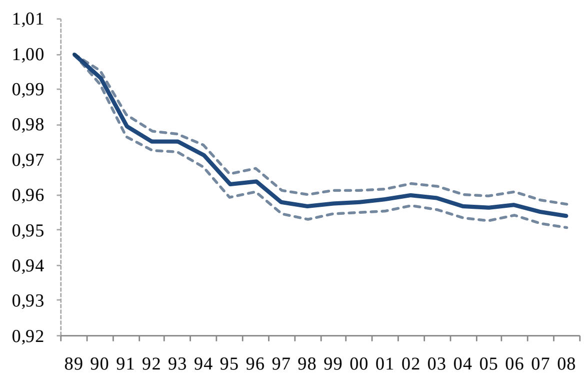
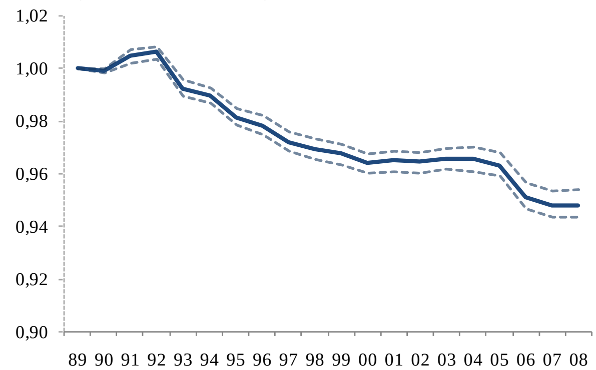


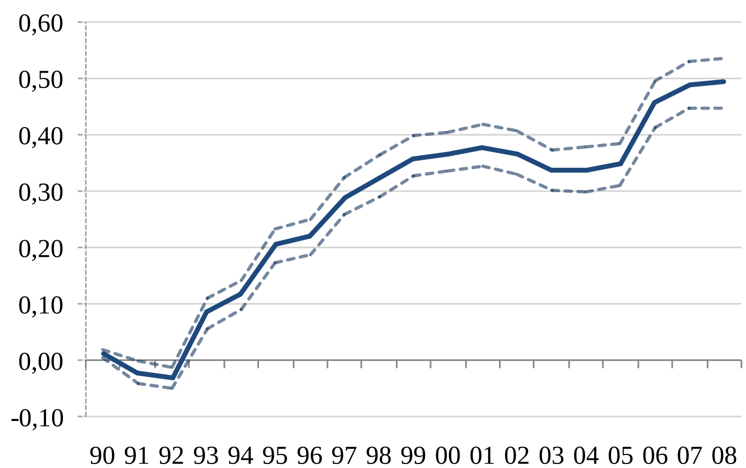
Figure 5. Intermediate goods chained lambda ratio



The last ingredients to the welfare calculation are the weights of imported consumption goods on aggregate consumption and the weights of imported intermediates on aggregate value added. The weight of imports on consumption is readily obtained from national account and import data, and turns out to be less than 2%. For intermediates, the only readily available production cost figure for the Brazilian economy, with intermediates discriminated as one of its components, is the annual industrial production research conducted by IBGE, the national statistics agency. This was used as a proxy for the economy wide ration. Given industrial production represents approximately one third of the economy wide value added, we were able to estimate the weight of imports on economy wide intermediate use as 8% to 10%. Note that this is a very inaccurate estimate, and one with important consequences, since it will tend to emphasize intermediate goods.

Under these assumptions, Figure 6 plots the accumulated gains from variety of imported goods in the Brazilian economy from 1989 to 2008. With regards to the dynamics of the welfare effects of changes in varieties imported, the gains from consumer goods in the beginning of the nineties were not strong enough to compensate for the opposing fall in intermediates, but did contribute to a strong positive trend during the remaining of the decade. The remaining dynamics, particularly the large welfare gains in 2006/07, mostly reflects intermediate goods. Summing up, the accumulated welfare gain in 2008 added up to 0.49%, with bootstrap confidence interval of (0.45%, 0.54%). In words, approximately a half percentage point of the economy's aggregate gross product would be traded for by consumers and firms only to gain access to new varieties of imported goods, supposing the new varieties under consideration were the ones that actually became available during the previous decade.

Figure 6. Gains from Variety 1990-2008
Equivalent Variation in terms of initial income



An important caveat is that welfare gains estimates refer to partial equilibrium effects, in the sense that domestic production reactions to increased foreign competition and possible welfare implications of this behavior are not accounted for in the estimates. The common wisdom in the literature is that domestic goods varieties would reduce after the trade liberalization and the introduction of cheaper varieties delivered by more efficient foreign firms; therefore, the partial equilibrium welfare gains may overstate general equilibrium gains. But theoretical arguments pertinent to the matter are not conclusive. Indeed, the present author has extended the workhorse model used in the literature to allow

for labor productivity effects which actually expand the availability of domestic varieties after trade liberalization - in that case, the gains from variety estimates from this paper define lower bounds on the general equilibrium gains.

6. Conclusion

The welfare calculations point to gains from increased variety of the order of 0.5% of the gross national product in the period from 1989 to 2008 in the Brazilian economy. The welfare gains was obtained taking into full account the strong heterogeneity in microeconomic elasticities of substitution parameters, as well as the different channels through which new varieties of consumer or intermediate goods affect the economy.

We document strong heterogeneity in elasticities of substitution and explored consequences of this heterogeneity for the applied researcher. First, aggregate elasticities of substitution for macroeconomic sectors and other broad economic categories were obtained as weighted averages of the underlying products. In support for this estimator, we stress the correlation of estimated aggregation bias and an above average presence of high elasticity goods in the economic sector, a correlation pattern required by the model that motivated the aggregate estimator. For the practitioner, this result point to an upward revision in aggregate elasticity estimates obtained from aggregate data and traditional time-series methods. Another consequence of heterogeneity and the proper account for product variety is to show that traditional import price indexes have accumulated an upward bias of up to 5% in the period under study, with important potential for biasing and confounding aggregate trade equation estimates of price and income effects.

The distinction of consumer and intermediate goods in the welfare model proved to be a particularly important contribution. Brazil is much closed when it comes to consumption than when it comes to the use of imported intermediate goods. Therefore, the productivity gains from intermediate goods have a greater impact on aggregate welfare. The distinction is also important to an adequate depiction of the dynamic gains from varieties, particularly when coupled with our chained index approach, since intermediate goods show a stark loss and gain in varieties in, respectively, the beginning of the nineties and the last years in the sample, while consumer goods show a marked increase in varieties only during the nineties.

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Domestic Varieties Under Free Trade: Shrinking or Expanding?

Abstract

The reduction in domestic varieties after trade liberalization, a result from trade models with heterogeneous firms [Melitz (2003)], is shown to be reverted as soon as we allow for a labor productivity effect. This is accomplished either through a complementary intermediate factor of production more widely available after trade, or through a technological upgrading effect from trade towards more productive skilled labor. The exact condition for an increase in domestic varieties of final consumption goods is for average firm revenue to increase less then proportionally to real wages. Conditions on primitives are also explained, along with implications for observable correlates, thus allowing inference on general equilibrium effects on domestic varieties production after exposure to foreign competition.

JEL Classification: D43, D60, F12, F16, L11, L13

Keywords: domestic variety; intermediate goods; skill premium; real wages; gains from trade; imperfect competition; firm heterogeneity.

1. Introduction

Expanding the variety of imported goods brings direct benefits to consumers; but the general equilibrium responses of the economy may also bring less clear-cut developments. Single sector trade models with heterogeneous firms suggest foreign competition compresses the diversity of domestic goods [Melitz (2003)]. Moreover, the measure of domestic producers displaced by foreign competitors is increasing in the average efficiency of foreign firms. Therefore, even in the case of a net increase in the variety of goods consumed, marginal imported goods are pricier than the domestic ones they displace, thus offsetting any gains from variety [Arkolakis, Demidova, Klenow and Rodriguez-Clare (2009)]. In light of these arguments, empirical efforts to estimate partial equilibrium gains from variety of imported goods [Broda and Weinstein (2006), Barroso (2009)] appear to be completely misplaced; instead, research should focus on the single force behind general equilibrium gains from trade, namely, the selection of more productive firms selling on average at better prices. In this paper, after minor changes to the Melitz model, we restore the relevance of the empirical estimates and qualify the reduction in domestic varieties. The crucial idea - with the hindsight that aggregate payroll is tightly related to the measure of domestic varieties -, is to increase labor productivity either by introducing endogenous intermediate goods or endogenous technological choice towards more skilled labor. The question, therefore, is if heterogeneous factors of production are sufficient to expand real payroll and domestic varieties while preserving the productivity gains from trade liberalization.

The possibility of an anti-variety effect is clearly stated in Melitz (2003) original paper, with high exporter fixed costs being the source of the problem. Indeed, higher fixed cost abroad implies greater selection and lower prices for imported varieties; as a result, imports displace more domestic competitors. Baldwin and Forslid (2004) develop the intuition further and show that total variety actually falls after trade whenever the fixed cost in the domestic market is higher for foreign competitors. In a recent paper, Arkolakis, Demidova, Klenow and Rodriguez-Clare (2009) extended the result for the case of many asymmetric countries. Moreover, they show gains from trade are independent of any variety effect, resting only on the dispersion parameter of the productivity distribution, that is, on

the magnitude of the firm selection effect; this strong result follows from the offsetting movement of import prices. The authors use their result to smoothly question the empirical gains from variety literature, including the exercise proposed as motivation for their study.

An important contribution from this paper is to incorporate Ethier (1982) and Romer (1993) approach of differentiated intermediate goods into the otherwise single sector Melitz model. Most of the analytical structure and intuition from the later is preserved in the process. For example, the free entry and the zero cutoff profit condition, a hallmark of the model which determine average profit and marginal entrant productivity in the final goods sector, now also apply to the intermediate goods sector. The economy is supposed to use a composite factor of production built from labor and intermediate goods through a constant returns to scale technology. The constant returns assumption allows one to link aggregate revenue and labor payroll just as in the Melitz model, but with real wage now endogenous and increasing in the aggregate quantity of intermediate goods. By the well know selection effect from trade, opening the economy raises productivity in the intermediate goods sector. Conditions for this selection to have a positive impact on intermediate goods production - which leads to a higher real wage and a higher measure of domestic varieties -, are studied both in general terms and in a parameterized economy. Somewhat ironically, additional welfare effects from the inclusion of a tradable intermediate goods sector follow only from the increased productivity in the sector, irrespective, therefore, of the greater or lesser availability of varieties of intermediate goods.

Another contribution is to show that not only intermediate goods but, as a general matter, heterogeneous factors of production may increase domestic varieties under free trade. To explore this issue we could consider a differentiated labor sector structurally similar to an intermediate sector. But since labor is non-tradable, it is not exposed to the selection effect from trade which was at the center of the argument before. An alternative solution is to have a discontinuous technology such that the opportunity to export leads more firms to demand skilled workers, which is exactly our approach. Of course, the choice of production technologies with different skill intensities is a frequent assumption in the trade literature, particularly for authors working in the skill premium issue. For example, Yeaple (2003) deduces endogenous firm heterogeneity and skill premium from

technological choices and a given distribution of workers among skill levels. The approach here is more closely related to Melitz in style, with entrants uncertain over the investment costs of upgrading to high technology. The model shows a positive relation between wage premiums for skilled workers and expansion of domestic goods varieties.

The intermediate goods and skilled labor extensions of the Melitz model point to useful alternatives to identify the general equilibrium responses of the economy to foreign competition. The additional structure imposed generates additional consequences, such as changes in the relative price of intermediate goods or changes in the skill premium. Therefore, the models bring more information to bear on the question of the response of domestic varieties to trade liberalization. The empirical literature following Broda and Weinstein (2006) have focused on the partial equilibrium effect of imported goods mostly under the assumption of a single sector. Barroso (2009) introduces an intermediate goods sector that allows for a productivity channel similar to this paper, but does not model imported goods interaction with the domestic production sector. The model developed here shows the estimated effect in the literature has a good chance to be a lower bound on general equilibrium welfare gains, and not a spurious artifact of the partial equilibrium approach, as suggested by single sector models. Moreover, the approach taken here means readily available evidence on relative prices and wages can be used to evaluate how likely the lower bound hypothesis proves to be. Finally, conditions for expanding domestic varieties are clearly stated as restrictions on elasticity and dispersion parameters emerging from selected parameterized examples.

2. Trade model with heterogeneous firms and an intermediate sector

2.1. Preferences and technology

Consider two identical economies without labor mobility between them. By symmetry, we make no distinction of variables from different countries. The economy consists of a final goods sector and an intermediate goods sector denoted respectively by $s = x, y$. Each sector has associated quantity and price indexes

(1) $$Q_s = \left(\int_{v \in \Omega_s} q_s(v)^{\rho_s} dv \right)^{\frac{1}{\rho_s}}, \qquad P_s = \left(\int_{v \in \Omega_s} p_s(v)^{1-\sigma_s} dv \right)^{\frac{1}{1-\sigma_s}}$$

where $\sigma_s = 1/(1-\rho_s)$ is the elasticity of substitution parameter. The quantity index Q_x represents consumer preferences over the continuum of available final goods; Q_y represents firm technology to bundle available varieties of intermediate goods into an aggregate input for further production stages. The available goods may include domestic and imported varieties. Whenever a distinction is necessary, variables will be superscripted with d or $*$ to denote domestic or international activity variables respectively.

Price indexes are the minimum cost to obtain a unit of the aggregate quantities choosing varieties $q_s(.)$ with given prices $p_s(.)$. $R_s = P_s Q_s$ is the required expenditure to obtain Q_s units of the aggregate quantity. Of course, expenditure is also the revenue for firms selling the relevant products. By symmetry, revenue from foreign firms in the domestic markets is the same as revenue from domestic firms abroad, thus ensuring balance of payments equilibrium and the equivalence of aggregate revenue and expenditures in each sector under free trade. From the cost minimization problems we obtain demand and expenditure functions

(2) $$q_s(v) = Q_s \left(\frac{p_s(v)}{P_s} \right)^{-\sigma_s}, \qquad \underbrace{r_s(v)}_{q_s(v)p_s(v)} = \underbrace{R_s}_{P_s Q_s} \left(\frac{p_s(v)}{P_s} \right)^{1-\sigma_s}$$

Consumers supply labor inelastically at the aggregate level L. Each variety is produced by a single firm. Labor and intermediate goods are bundled by firms in a second production stage to obtain a composite factor $Z = Z(L, Q_y)$ using a constant returns to scale technology, so that firms may be thought to coordinate on the aggregate decision. Given the price of labor w_l and of intermediate goods P_y, the minimum cost to assemble and therefore the price to obtain the composite factor is denoted by w_z. Firm productivity φ determines the required quantity of the composite factor for a firm to obtain a particular output

$$z_s^d(\varphi) = \frac{q_s^d(\varphi)}{\varphi} + f_s = \frac{r_s^d(\varphi)}{\varphi p_s^d(\varphi)} + f_s \tag{3}$$

where f_s is the fixed cost in units of the composite factor.

If the firms decides to serve foreign markets by exporting its output, we set the indicator $\chi(\varphi) = 1$ and note that the firm faces a fixed cost f_s^*; otherwise we set $\chi(\varphi) = 0$. In the case of a closed economy, χ is constrained to be zero. Export is also subject to an iceberg transportation cost $\tau > 1$, so that the actual marginal cost is τ/φ. As in (3), the exporter marginal and fixed costs define $z_s^*(\varphi)$ required by the production level.

Before starting production firms must sunk an entry cost f_s^e without information on their productivity level, which is only known to have distribution $\mu(\varphi)$ on the support $(0, \infty)$ with a continuous distribution function. Upon entry, firms learn their productivity and decide if profit levels are sufficient to cover the relevant fixed costs of their choices. Firms optimize profits given the demand structure in (2), thus leading to the following pricing and profit relations:

$$p_s^d(\varphi) = \frac{1}{\rho_s}\frac{w_z}{\varphi}, \qquad p_s^*(\varphi) = \tau p_s^d(\varphi) \tag{4}$$

$$\pi_s(\varphi) = \underbrace{r_s^d(\varphi) - z_s^d(\varphi) w_z}_{\pi_s^d(\varphi)} + \chi(\varphi)\underbrace{\left[r_s^*(\varphi) - z_s^*(\varphi) w_z\right]}_{\pi_s^*(\varphi)} \tag{5}$$

Since only firms with positive profit would choose positive production levels, only firms with productivity draws above some minimum cutoff level ϕ_s would be successful entrants. These firms are able to charge low enough prices to capture a relevant fraction of the sectors' revenue while also saving on production costs. Given the higher fixed costs of the export activity, only firms with productivities at least above an even higher cutoff $\phi_s^* > \phi_s$ would choose to become exporters. Since productivity is unbounded, cutoff productivities for entrants and exporters are well defined. In the case of a closed economy we impose $\phi_s^* = \infty$.

Each period, incumbent firms face an exogenous probability δ of exiting the market. As a result, the expected profit stream conditional on successful entry is $\sum_t (1-\delta)^t \bar{\pi}_s = \bar{\pi}_s/\delta$ where $\bar{\pi}_s = E_{\varphi_s > \phi_s} \pi_s(\varphi_s)$ is the average incumbent profit. Moreover, the ex-ante expected profit stream is $\mu(\varphi_s > \phi_s)\bar{\pi}_s/\delta$, which is the crucial variable in firm entry decision.

As a matter of notation, define average productivity $\tilde{\varphi}_{\phi_s} = \left(E_{\varphi_s > \phi_s} \varphi^{\sigma-1}\right)^{1/(\sigma-1)}$. Thus, average incumbent profit may be written $\bar{\pi}_s = \pi_s^d\left(\tilde{\varphi}_{\phi_s}\right) + \mu(\varphi_s > \phi_s^*)\pi_s^*\left(\tilde{\varphi}_{\phi_s^*}\right)$. Similarly, with the appropriate average notion, aggregate price index becomes

(6) $$P_s = (M_s^d)^{\frac{1}{1-\sigma_s}} \bar{p}_s = (M_s^d)^{\frac{1}{1-\sigma_s}} \left[p_s^d\left(\tilde{\varphi}_{\phi_s}\right)^{1-\sigma_s} + \mu(\varphi_s > \phi_s^*) p_s^*\left(\tilde{\varphi}_{\phi_s^*}\right)^{1-\sigma_s}\right]^{\frac{1}{1-\sigma_s}}$$

where M_s^d is the measure of domestic firms and therefore of domestically produced varieties. Alternatively, define a trade adjusted average productivity

(7) $$\tilde{\varphi}_s^t = \left[\frac{M_s^d}{M_s}\left(\tilde{\varphi}_{\phi_s}\right)^{\sigma_s - 1} + \frac{M_s^*}{M_s}\left(\tilde{\varphi}_{\phi_s^*}/\tau\right)^{\sigma_s - 1}\right]^{\frac{1}{\sigma_s - 1}}$$

where $M_s = M_s^d + M_s^* = M_s^d + \mu(\varphi_s > \phi_s^*) M_s^d$ is the measure of varieties in the domestic market including imported ones; using the definition,

(8) $$P_s = (M_s)^{\frac{1}{1-\sigma_s}} p_s^d(\tilde{\varphi}_s^t)$$

Analogous representations may be deduced for aggregate quantities and revenues. As the notation makes clear, marginal entrant productivity is the single determinant of all aggregate variables; and that is also the case for open economies for

(9) $$\left(\frac{\phi_s^*/\tau}{\phi_s}\right)^{\sigma_s-1} = \frac{r_s^*(\phi_s^*)}{r_s^d(\phi_s)} = \frac{f_s^*}{f_s}$$

allows one to write marginal exporter productivity in terms of the marginal entrant.

2.2. Steady state equilibrium

Free entry drives the ex-ante expected profit to the sunk cost of entry, that is

(FE) $$\bar{\pi}_s = \delta f_s^e/\mu(\varphi_s > \phi_s)$$

Average incumbent profit is determined by the zero cutoff profit condition[9]

(ZCP) $$\begin{aligned}\bar{\pi}_s &= \pi_s^d\left(\tilde{\varphi}_{\phi_s}\right) + \mu(\varphi_s > \phi_s^*)\pi_s^*\left(\tilde{\varphi}_{\phi_s^*}\right) \\ &= k_s(\phi_s)f_s + \mu\left(\varphi_s > \phi_s^*(\phi_s)\right)k_s\left(\phi_s^*(\phi_s)\right)f_s^*\end{aligned}$$

where $k_s(\phi_s) = \left(\left(\tilde{\varphi}_{\phi_s}/\phi_s\right)^{\sigma_s-1} - 1\right)$ and w_z is normalized to one.

Intersection of both conditions defines equilibrium average profit and cutoff productivity for each sector; existence is established in Melitz (2003). Trade liberalization shifts the zero cutoff profit condition up because the second term turns positive for some productivity levels. Note this condition is decreasing in productivity while free entry is increasing. As a result, equilibrium average profit and cutoff productivity are higher than in autarky, which amounts to the selection effect from trade.

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[9] $\pi_s\left(\tilde{\varphi}_{\phi_s}\right) = \frac{r_s\left(\tilde{\varphi}_{\phi_s}\right)}{\sigma_s} - f_s w_z = \frac{1}{\sigma_s}\left(\frac{\tilde{\varphi}_{\phi_s}}{\phi_s}\right)^{\sigma_s-1} r_s\left(\phi_s\right) - f_s w_z = \frac{1}{\sigma_s}\left(\frac{\tilde{\varphi}_{\phi_s}}{\phi_s}\right)^{\sigma_s-1}\sigma_s f_s w_z - f_s w_z = \left(\left(\tilde{\varphi}_{\phi_s}/\phi_s\right)^{\sigma_s-1} - 1\right) f_s w_z$, with the third equality following from zero profit at the marginal entrant.

An important feature of the model is that cutoffs are determined independently of the measure of firms (or varieties). To define the equilibrium measure we need to explore economy wide resource constraints and aggregate factor payments, which is our next step.

Cost minimization in the second production stage equalizes marginal productivities to the real factor prices in terms of the composite factor

$$Z_L(L, Q_y) = w_l, \qquad Z_y(L, Q_y) = P_y \tag{10}$$

By Euler theorem, factor payments exhaust output's value

$$w_l L + P_y Q_y = Z \tag{11}$$

Integrating (5) and adding both sector equations we obtain the aggregate composite factor used in production $Z^p = \mathrm{R} - \Pi$. Steady state requires successful entrants in each sector to equal the measure of exiting firms $\mu(\varphi_s > \phi_s)M_s^e = \delta M_s^d$. Therefore, composite factor used up in the entry process is $Z_s^e = M_s^e f_s^e = M_s^d \delta f_s^e / \mu(\varphi_s > \phi_s) = M_s^d \bar{\pi}_s = \Pi_s$, and adding across sectors $Z^e = \Pi$. Therefore, total composite factor amounts to

$$Z = R = R_x + R_y = P_x Q_x + P_y Q_y \tag{12}$$

From (11) and (12) we conclude $w_l L = R_x = M_x^d \bar{r}_x$. Note that $\bar{r}_x = \sigma_x(\bar{\pi}_x^d + f_x) + \mu(\varphi_s > \phi_s^*)\sigma_x(\bar{\pi}_x^* + f_x^*)$ is already determined by the FE and ZCP conditions. Using (9) we get the labor resource constraint

$$M_x^d \bar{r}_x = Z_L(L, Q_y) L \tag{LR}$$

Similarly, from (10-12) and the two alternative representations for R_y from the last section, we obtain a constraint for intermediate goods

$$M_y r_y(\tilde{\varphi}_s^t) = M_y^d \bar{r}_y = Z_Y(L, Q_y) Q_y \tag{13}$$

Using (8) to substitute M_y and then (10) to substitute for the price index, there follows

(14) $$\left(\rho_y \tilde{\varphi}_y^t\right)^{1-\sigma_y} r_y(\tilde{\varphi}_s^t) = Z_y\left(L, Q_y\right)^{\sigma_y} Q_y$$

From (2) and zero profit for the marginal entrant we have $r_y\left(\tilde{\varphi}_y^t\right) = \left(\tilde{\varphi}_y^t/\phi_s\right)^{\sigma_y - 1}\sigma_y f_y$. Substituting into (14) we obtain the intermediate goods resource constraint

(IR) $$h\left(\phi_y\right) = f\left(Q_y\right)$$

where $h\left(\phi_y\right) := \left({\rho_y}^{1-\sigma_y}\sigma_y f_y\right)\left(\phi_y\right)^{1-\sigma_y}$ and $f\left(Q_y\right) := Z_y\left(L, Q_y\right)^{\sigma_y} Q_y$. For some comparative static results, we will also use the equivalent formulation

(IR) $$g\left(\phi_y, Q_y\right) = z\left(Q_y\right)$$

where $g\left(\phi_y, Q_y\right) := h\left(\phi_y\right)^{-1} Q_y$ and $z\left(Q_y\right) := Z_y\left(L, Q_y\right)^{-\sigma_y}$.

The (IR) condition determines equilibrium Q_y. All other endogenous variables are thereafter determined. In particular, equilibrium measure of domestic varieties follows from (LR), prices for final goods from (6) and real wages from (10); prices and wages determine per capita welfare which is $W = w_l/P_x$.

For existence, uniqueness and comparative statics results, it is necessary to impose some assumptions. The following regularity assumptions are sufficient for our purposes.

(A1) Z is C^2 with $Z_l > 0$, $Z_y > 0$, $Z_{yy} < 0$, $Z_{yl} < 0$

(A2) $\lim_{q\to 0} f(q) = \infty$ and $\lim_{q\to\infty} f(q) = 0$

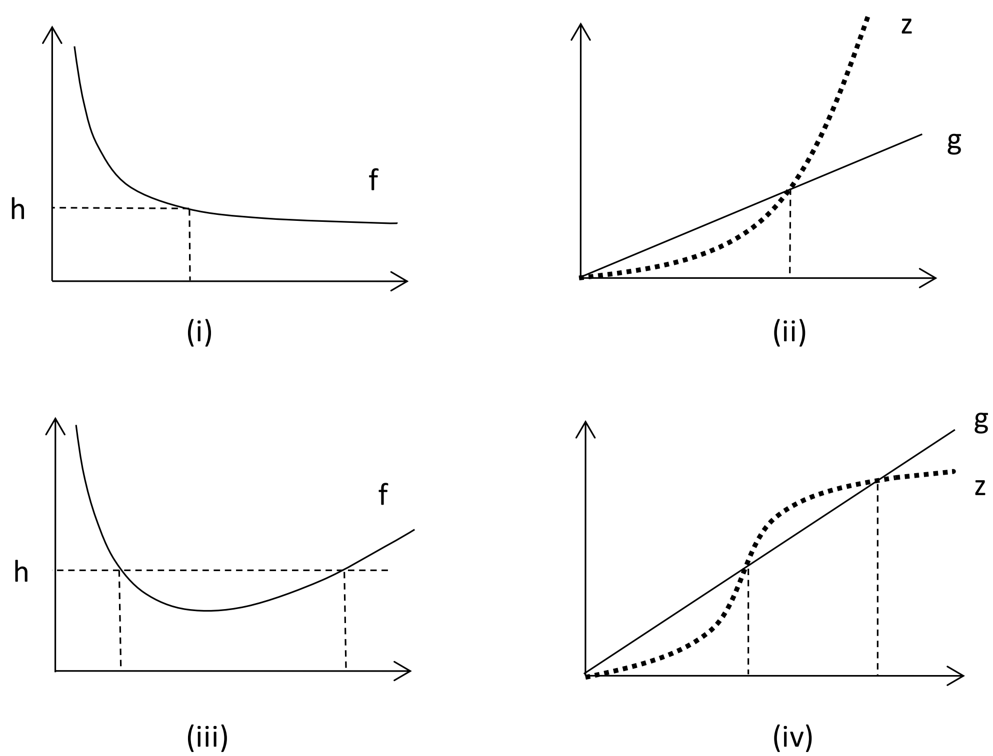
(A3) $\sigma_y \left|\varepsilon_y^{Z_y}\right| > 1$ for all q > 0, where $\varepsilon_y^{Z_y} = \frac{Z_{yy}(l,q)q}{Z_y(l,q)}$,

The CES assumption for Z is not consistent with (A2) and (A3). We develop results for the CES case in the examples bellow. Figure 1 illustrates both the regular and the CES cases.

Proposition 1. Suppose (A1-A2). Then there is $Q_y > 0$ that solves (IR). Additionally, suppose (A3). Than Q_y is unique.

Example. For a CES production function, $f(q) = (L^{\rho_z} + q^{\rho_z})^{\frac{1}{\rho_z}\cdot\frac{\sigma_y}{\sigma_z}} q^{1-\frac{\sigma_y}{\sigma_z}}$. Under the assumption $\sigma_y > \sigma_z$ the limit conditions (A3) run backwards, but a similar existence proof is obtained. As it turns out, the case $\sigma_y > \sigma_z$ has more desirable comparative statics properties - the assumption means it must be harder to substitute intermediate goods for labor than for other varieties of intermediate goods themselves. We focus on this case, illustrated in the lower panel of Figure 1. It is clear the model may have multiple equilibria. The equilibrium with lower level of intermediate goods resembles the regular case, while the one with high level of intermediate goods resembles the $\sigma_z > \sigma_y$ case. Existence may be a problem, but can be obtained for low levels of the labor force or high levels of entry costs in intermediate goods production.

Figure 1. Regular case (i, ii) and CES case (iii, iv)



For comparative statics, we can either work with the f representation or with the g and z representation. The crucial point is assumption (A3), which leads to a decreasing schedule for f(.). The analogous assumption in terms of z(.) would be convexity. In the following propositions, all regularity conditions (A1-A3) are admitted.

Proposition 2. Intermediate goods production is increasing in the labor force.

The intuition is that labor shifts down the z curve in Figure 1, panel (ii). Therefore it increases intermediate goods. The effect on other endogenous variables is ambiguous, since more labor and intermediate goods have opposing effects on real wages.

Proposition 3. Trade liberalization increases intermediate goods production, real wages and final goods sector revenues. The measure of domestic varieties increases if and only if revenues increase less than real wages.

Indeed, from Melitz, trade liberalization shifts the zero cutoff profit condition up in both sectors. This shifts the g curve up, thus increasing intermediate goods aggregate. Real wage increases due to labor complementarity with intermediate goods. From the (LR) condition, the effect on the measure of domestic varieties of final goods may be positive, which depends on the average revenues in the final goods sector raising less than real wages. The example bellow provides further intuition.

Example. From the (IR) condition it follows that $\varepsilon := \frac{dlnQ_y}{dln\phi_y} = \frac{1-\sigma_y}{1-\sigma_y\left|\varepsilon_y^{Z_y}\right|}$, where $\varepsilon_y^{Z_y}$ is the elasticity of intermediates marginal product to an increase in this input. Consider again the CES production function, where $\left|\varepsilon_y^{Z_y}\right| = \frac{L^{\rho_z}}{L^{\rho_z}+{Q_y}^{\rho_z}}\frac{1}{\sigma_z}$. In this case, ε is positive only for low values of intermediate goods production and only if $\sigma_y > \sigma_z$. The relative magnitude of revenue and wage effects depends on distribution assumptions on firm productivities. Suppose distributions are Pareto with dispersion parameter θ so that $k_s(\phi_s) = [\theta/(\theta-\sigma_s+1)]^{\sigma_s-1}-1$. Taking the difference of the logarithm of the (LR) condition for a closed and open economy, after some algebra, one may show that domestic varieties increase after trade liberalization if and only if $\varepsilon_y^{Z_L}\varepsilon f > \theta$, where f > 0 depends on the fixed

costs in the model. Intuitively, intermediate goods must have strong complementarity with labor and mildly decreasing marginal product, while the selection effect must be kept bounded by a low dispersion parameter.

The last proposition is the main result from the model, as it establishes a clear possibility for expanding domestic varieties in comparison with the unequivocal shrinking results from the previous literature. The example is also important. It makes the intuitive point that expanding domestic varieties requires the complementarity effects emphasized in this paper to be large relative to the usual selection effects, which is a strictly empirical matter summarized by well-defined elasticity and dispersion parameters.

To close this section, we study the welfare impact of trade. Per capita welfare is simply the aggregate quantity index on final goods divided by the labor force,

$$W = \frac{R_x/L}{P_x} = w_l M_x^{\frac{1}{\sigma_s - 1}} \rho_x \tilde{\varphi}_x^t \tag{15}$$

Similarly to (9), we have

$$\left(\frac{\tilde{\varphi}_x^t}{\phi_x}\right)^{\sigma_s - 1} = \frac{r_x^d(\tilde{\varphi}_x^t)}{r_x^d(\phi_x)} = \frac{R_x/M_x}{f_x} \tag{16}$$

Solving (16) for $\tilde{\varphi}_x^t$ and substituting in (15),

$$W = w_l M_x^{\frac{1}{\sigma_s - 1}} \rho_x \left(\frac{R_x/M_x}{f_x}\right)^{\frac{1}{\sigma_s - 1}} \phi_x = w_l \rho_x \left(\frac{M_x^d \bar{r}_x}{f_x}\right)^{\frac{1}{\sigma_s - 1}} \phi_x \tag{17}$$

From (17) and the comparative static results above, we obtain the following proposition:

Proposition 4. Trade liberalization has a selection effect ($\Delta\phi_x > 0$), a real wage effect ($\Delta w_l > 0$) and a payroll effect ($\Delta R_x > 0$), all of which contribute positively to welfare. The payroll effect may be further decomposed into an final good average expenditure effect ($\Delta\bar{r}_x > 0$) and a final good domestic variety effect (ΔM_x^d), with positive welfare consequences from average expenditure and a possible positive one from domestic varieties, conditional on its measure increasing after trade.

In the above decomposition, only the selection effect is already in the Melitz model. Complementary between intermediates and labor is responsible for the wage effect. It also decouples payroll from the labor force, thus allowing for the average expenditure and domestic variety effects.

3. Trade model with heterogeneous firms and skill premium

3.1. Preferences and technology

Two identical economies are each endowed with a measure L_l of low skill labor and a measure L_h of high skill labor, both supplied inelastically to firms. Low skill labor is the numeraire and w denotes high skill labor wage (or skill premium).There is a representative household with preferences over final goods defined by the same quantity and price indexes from the last section. Expenditure minimization leads once again to demand and expenditure decisions (2) – without the sector sub-index - for each variety of the consumption good.

Each active firm produces a single variety and faces a fixed production cost f in units of low skill labor. There is a common knowledge low technology which uses a unit of low skill labor to obtain a unit of some consumer good variety. There is also a high technology process which uses a unit of skilled labor to get $\varphi > 1$ units of final goods of a particular variety; but firms face an investment cost of $1/\omega$ in low skill labor to access this technology. Before starting production firms must sunk an entry cost f^e without information on their investment efficiency ω, which has an atomless distribution $\mu(\omega)$ on the $(0, \infty)$ support. The exporting activity has a fixed cost f_* in units of low skill labor and is subject to an iceberg transportation cost $\tau > 1$. Upon entry, firms learn their efficiency, choose their technology and optimize profits along demand curves, with decisions as in (4) and (5) from the previous section stated for the adequate factor of production.

With a probability δ firms exit the market and we define the expected discounted profit as before; in this case, it is just discounted average incumbent profit, since all firms can produce at the common knowledge technology at non-negative profits (otherwise there would be no demand for low skill labor). Low tech profits must be bounded above by δf^e, otherwise there would be an unbounded measure of entrants. We assume fixed export cost

f_* is high enough, so that only high tech firms can sustain exporting activity in equilibrium. Since $\pi_{l*} = \tau^{1-\sigma}(\pi_l + f) - f_* \leq \tau^{1-\sigma}(\delta f^e + f) - f_*$, this can be assured by $\tau^{1-\sigma}(\delta f^e + f) < f_*$. We also assume that the fixed export cost is not so high as to drive exporting profit below zero. If this conflicts with the first assumption, just impose technological upgrade as a necessary condition to export.

3.2. Steady state equilibrium

Let the technological upgrading probability in a closed economy be denoted by[10]

$$\mu(w, r_l) = \mu\left(\frac{1}{\omega} < \delta^{-1}(\pi_h - \pi_l)\right) = \mu\left(\frac{1}{\omega} < \frac{1}{\delta\sigma}(r_h - r_l)\right)$$
$$= \mu\left(\omega > \frac{\delta\sigma}{r_l[(\varphi/w)^{\sigma-1} - 1]}\right)$$

and the analogous probability for a open economy by

$$\mu_*(w, r_l) = \mu\left(\frac{1}{\omega} < \delta^{-1}(\pi_h - \pi_l + \pi_{h*})\right) = \mu\left(\frac{1}{\omega} < \frac{\delta}{\sigma}(r_h - r_l + r_{h*}) - f_*\right)$$
$$= \mu\left(\omega > \frac{\delta\sigma}{r_l[(\varphi/w)^{\sigma-1} - 1 + \tau^{1-\sigma}(\varphi/w)^{\sigma-1}] - \sigma f_*}\right)$$

Note these are decreasing functions of w and increasing functions of r_l. Moreover, we have $\mu_* > \mu \Leftrightarrow \pi_{h*} > 0$, a condition we maintain bellow, restricting the analysis to skill premiums below a break-even level for exporters. For ease of reference, let also $\omega(w, r_l)$ and $\omega_*(w, r_l)$ be the minimum cutoff efficiency levels associated with those probabilities. Then define the expected investment costs conditional on upgrading as

$$f^u(w, r_l) = E_{\omega > \omega_*(w, r_l)}(1/\omega)$$

[10] In this section, we will use a lot $\pi_s = r_s/\sigma - f$ and $\pi_h - \pi_l = \frac{1}{\sigma}\left(\frac{\varphi}{w}^{\sigma-1} - 1\right) r_l$ which follow from the last sections equations (2)-(5) and this section technology.

$$f_*^u(w, r_l) = E_{\omega > \omega_*(w, r_l)}(1/\omega)$$

for the closed and open economy cases, respectively. The functions are decreasing in w and increasing in r_l, with $f_*^u > f^u$ for the same upward shift condition $\pi_{h*} > 0$.

The free entry condition states expected discounted profits must equal expected total investment before production. For a closed and open economy the conditions are

(FE) $$(\pi_l)_{(r_l)} + \mu(w, r_l)(\pi_h - \pi_l)_{(w, r_l)} = \delta f^e + \delta \mu(w, r_l) f^u(w, r_l)$$

(FE*) $$(\pi_l)_{(r_l)} + \mu_*(w, r_l)(\pi_h - \pi_l + \pi_{h*})_{(w, r_l)} = \delta f^e + \mu_*(w, r_l)\delta f_*^u(w, r_l)$$

where we emphasize profits are functions of skill premium and low tech revenues.

Lemma 1. Suppose firms expect to upgrade after entry, so that $\pi_h - \pi_l > \delta f^u(w, r_l)$ and $(\pi_h - \pi_l + \pi_{h*}) > \delta f_*^u(w, r_l)$. Suppose decreasing the skill premium w relaxes these inequalities, with technological premium increasing faster than the expected upgrading costs. Then, each free entry conditions establish a monotone *increasing* relation $r_l(w)$. Moreover, if export profit is expected to more than compensate the increase in average upgrading costs after trade liberalization, the free entry curve shifts down once liberalization occurs.

Example. Consider firms' efficiency in upgrading is drawn from a Pareto distribution with dispersion parameter θ and minimum b. Then $\mu(w, r_l)\big((\pi_h - \pi_l) - \delta f^u(w, r_l)\big) = \left(\frac{b}{\delta}\right)^\theta (\pi_h - \pi_l)^\theta \left[(\pi_h - \pi_l) - \frac{\theta}{\theta+1}(\pi_h - \pi_l)\right]$. Therefore, the free entry condition for a closed economy is

$$\left(\frac{r_l}{\sigma} - f\right) + \left(\frac{b}{\delta}\right)^\theta \left\{\frac{1}{\sigma}\left(\left(\frac{\varphi}{w}\right)^{\sigma-1} - 1\right) r_l\right\}^{\theta+1} \left(1 - \frac{\theta}{\theta+1}\right) = \delta f^e$$

Note that a Pareto distribution satisfies the condition in the Lemma for an increasing relation $r_l(w)$. In the open economy, the term in brackets changes to

$$\left\{\frac{1}{\sigma}\left(\left(\frac{\varphi}{w}\right)^{\sigma-1}-1\right)r_l+\pi_{h*}\right\}$$

Therefore, supposing positive export profits, the $r_l(w)$ curve shifts down.

Denote ${L_l}^p$ and ${L_h}^p$ low skill and high skill labor demand for production in their respective sectors. Let ${L_l}^{p,h}$ denote low skill labor demand due to fixed cost in the high tech firms. We have

$$\begin{gathered}{L_l}^p = R_l - \Pi_l + {L_l}^{p,h}\\ {L_h}^p w = R_h - \Pi_h - {L_l}^{p,h}\end{gathered}$$

where sub-indexes on total revenue and profits refer to the technology of the aggregated firms. The measure of entrants is M^e, and incumbent domestic firms with low M_l^d and high M_h^d technology amount to $M^d = M_l^d + M_h^d$ firms. Steady state requires

$$\begin{gathered}{L_l}^{e,l} = (1-\mu)M^e f^e = \delta M_l^d f^e\\ {L_l}^{e,h} = \mu M^e(f^e + f^u) = \delta M_h^d(f^e + f^u)\end{gathered}$$

where ${L_l}^{e,l}$ and ${L_l}^{e,h}$ refer to low skill labor demanded for investment purposes. Therefore

$${L_l}^e = {L_l}^{e,l} + {L_l}^{e,h} = M^d(\delta f^e + \delta\mu f^u) = M^d E\pi = \Pi$$

Adding up labor demand for all purposes and equating with the available measure,

$$\begin{gathered}L_l = \mathrm{R_l} - \Pi_l + \Pi + {L_l}^{p,h}\\ L_h w_h = \mathrm{R_h} - \Pi_\mathrm{h} - {L_l}^{p,h}\end{gathered}$$

Therefore, we get the resource balance condition

(RB) $$L_l + L_h w_h = \mathrm{R} = M^d E r$$

Of course, the same condition holds in the open economy by the exact same argument. In a sense, this is a central result in the model, since payroll depends not only on the labor force, but also on skill distribution and skill premium, opening a possible means for expanding domestic varieties.

Consider now the equality of high skill labor supply and demand. For a closed economy, we have

$$L_h = \mu M^d \frac{1}{\varphi} q_h = \mu M^d P^{\sigma-1} R \rho^{\sigma} \frac{1}{w}\left(\frac{\varphi}{w}\right)^{\sigma-1}$$

Using resource balance (RB) and the price index definition

$$P^{\sigma-1} R = \frac{(L_l + L_h w)}{M^d \frac{1}{\rho}\left(1 - \mu + \mu\left(\frac{\varphi}{w}\right)^{\sigma-1}\right)}$$

Substituting and rearranging, we obtain the skilled labor condition

(SL) $$\frac{1}{(\varphi/w)^{\sigma-1}}\left(\frac{1}{\mu(w, r_l)} - 1\right) + 1 = \frac{L_l}{L_h}\frac{1}{w} + \rho^{\sigma+1}$$

For the open economy we replace q_h with $(1+\tau^{-\sigma})q_h$ and add the term $\mu\tau^{1-\sigma}(\varphi/w)^{\sigma-1}$ to the price index, resulting in the skilled labor condition for an open economy

(SL*) $$\frac{1}{(\varphi/w)^{\sigma-1}}\left(\frac{1}{\mu_*(w, r_l)} - 1\right) + 1 + \frac{1}{\tau^{\sigma-1}} = \frac{L_l}{L_h}\frac{1}{w} + \rho^{\sigma+1} + \frac{1}{\tau^{\sigma}}$$

Lemma 2. The skilled labor condition defines a monotone increasing relation $r_l(w)$ between low skill revenue and skill premium. Opening up the economy, on the one hand, tends to reduce the price index, thus reducing each firm local demand (last term on the left hand side); on the other hand it stimulates technological upgrading ($\mu_* > \mu$) and the production of high tech firms now serving foreign markets (last term on the right hand side). Given $\tau^{-\sigma} < \tau^{1-\sigma}$, the combined effect of the last terms in both sides of the equation is to shift the curve up relative to a closed economy. On the opposite direction, the increased incentive for technological upgrading shifts the curve down.

Example. Lets continue the Pareto distribution example. We can solve for $1/\mu$ the skilled labor conditions and substitute the probabilities implied by the Pareto assumption; then we solve for r_l. For the closed economy we have

$$\left(\left(\frac{\varphi}{w}\right)^{\sigma-1} - 1\right) r_l = \left\{\left(\frac{b}{\delta\sigma}\right)\left[\left(\frac{\varphi}{w}\right)^{\sigma-1}\left(\frac{L_l}{L_h}\frac{1}{w} + \rho^{\sigma+1} - 1\right) + 1\right]^{1/\theta}\right\}^{-1}$$

while for the open economy we have

$$\left(\left(\frac{\varphi}{w}\right)^{\sigma-1} - 1\right) r_l + \pi_{h*} = \left\{\left(\frac{b}{\delta\sigma}\right)\left[\left(\frac{\varphi}{w}\right)^{\sigma-1}\left(\frac{L_l}{L_h}\frac{1}{w} + \rho^{\sigma+1} - 1 + \frac{1}{\tau^{\sigma}} - \frac{1}{\tau^{\sigma-1}}\right) + 1\right]^{1/\theta}\right\}^{-1}$$

For low levels of π_{h*} the increase in the right hand side after trade dominates and the curve shifts up. For sufficiently high export profit levels (low fixed costs, high productivity) the higher incentive for technological upgrading dominates and the curve shifts down.

Proposition 5. Maintain the assumptions stated in the Lemmas 1 and 2 for the free-entry and skilled labor curves to be increasing. Suppose the derivative of the skilled labor curve to be smaller than the derivative for the free-entry curve at $w = 0$. Then there is an equilibrium.

Example. Continuing our example the derivative at zero of the skilled labor curve is zero while the same derivative for the free-entry curve is infinite. Equilibrium exists by the previous proposition.

Proposition 6. Maintain the assumptions stated in the Lemmas 1 and 2 for the free-entry and skilled labor curves to be increasing *and* to shift down with trade liberalization - sufficient export profitability is the key necessary condition. Then opening up to trade increases the skill premium but has ambiguous effects on low tech revenue. Consider the resource balance condition; the effect on the measure of domestic varieties of consumer goods may be positive, which depends on the average revenues in the final goods sector rising less than the skill premium.

This is the main result for this model, for it again establishes the possibility of expanding domestic varieties. The main idea shared by this result and Proposition 3 from the model with intermediate goods, is to explore the extra degree of freedom in factor markets to increase factor payments, increasing aggregate expenditure and firm entry. The representative consumer framework makes welfare comparisons less compelling in the present context, since endowment distribution was not discussed at any point. But it is clear per capita welfare is increased through higher wages and increased varieties.

4. Conclusion

Trade models with heterogeneous firms have emphasized welfare gains from the selection effect from trade liberalization, an effect so strong it could compensate the necessary reduction in domestic varieties or even in overall consumption goods varieties. This paper makes important qualifications to this conventional wisdom. Indeed, all the previous results may be reverted as soon as we allow for an increase in labor productivity following trade liberalization - either through a complementary intermediate factor of production more widely available after trade or through a technological upgrading effect from trade towards more productive skilled labor. The exact condition for an increase in domestic varieties of final consumption goods is for average firm revenue to increase less

then proportionally to real wages. The two models, intermediate goods and skilled labor, use completely different mechanisms to obtain this exact same result, and thus illustrate a more general link from productivity and gains from variety. Firm heterogeneity is explored in an essential manner in both mechanisms, particularly in the skill premium model, where firms choose endogenously to use different technologies. For the empirical literature, the models offer important insights and incentives in the form of correlates for domestic variety increase (such as higher skill premium or lower price index for intermediary goods) and primitive parameters associated with the results.

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Appendix.

Lemma. Under (A1), f is strictly decreasing if and only if (A3).

Proof. $f' = \sigma_y Z_y(l,q)^{\sigma_y - 1} Z_{yy}(l,q)q + Z_y(l,q)^{\sigma_y} = Z_y(l,q)^{\sigma_y}\left(1 - \sigma_y \left|\varepsilon_y^{Z_y}\right|\right)$.

Proof of Proposition 1. From the limit at zero condition, there is a $q_a > 0$ such that $f(q_a) > h > 0$. Similarly, from the limit at infinity condition, there is a $q_b > q_a$ such that $f(q_b) < h$. Continuity for f follows from continuity assumption for Z. From the intermediate value theorem, there is a $q^* \in [q_a, q_b]$ such that $f(q^*) = h$. Moreover, if the elasticity condition holds, uniqueness follows from the above lemma.

Proof of Proposition 2. Since $df/dl = \sigma_y Z_y(l,q)^{\sigma_y - 1} Z_{yl}(l,q)q$>0, an increase in l shifts up the f schedule. From the above lemma, there must be an increase in q to restore the (IR) condition. These changes have ambiguous effects on Z_l and Z_y.

Proof of Proposition 3. From Melitz, ϕ_y increases after trade. From the above lemma, q must increase to restore the (IR) condition.

Proof of Proposition 5. Taking the limit of the respective conditions when w goes to infinity, we get $r_l/\sigma - f = \delta f^e$ for free-entry and $\mu(\infty, r_l) = 1$ for skilled labor, so that the first is bounded above while the second is unbounded. When w goes to zero, we now get $\infty r_l/\sigma = \delta f^e + \delta \bar{f}^u + f$ for free-entry and $1 = \infty\mu(0, r_l)$ for skilled labor; and r_l goes to zero in both curves. Therefore existence requires the derivative of the skilled labor curve to be smaller than the derivative for the free-entry curve at $w = 0$.