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**Working Paper**

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LICOS Discussion Paper, No. 187

**Provided in Cooperation with:**

LICOS Centre for Institutions and Economic Performance, KU Leuven

*Suggested Citation:* Olper, Alessandro; Raimondi, Valentina (2007) : Market access asymmetry in food trade, LICOS Discussion Paper, No. 187, Katholieke Universiteit Leuven, LICOS Centre for Institutions and Economic Performance, Leuven

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## LICOS Discussion Paper Series

Discussion Paper 187/2007

### Market Access Asymmetry in Food Trade

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# Market access asymmetry in food trade

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[This version: May 2007]

## Abstract

Using a bilateral trade equation derived from a monopolistic competition model, we investigated market access reciprocity in food trade among the US, Canada, the EU and Japan. We explore country and industry-specific market access asymmetry through the border effect approach, re-challenging the underlying main explanations. Our findings reveal marked asymmetry in reciprocal trade openness; indeed, access to the food markets of the US and Japan appears significantly easier than reciprocal access to both Canada and, especially, the EU. Policy trade barriers, firstly in the forms of NTBs, the degree of product differentiation and 'home bias' in preferences, are all important factors in explaining border effects. Moreover, several stylized facts suggest that border effect interpretation should also be based on political economy arguments.

**JEL Classification:** F13, F14, Q17

**Keywords:** gravity, market access, asymmetry, food trade, NTBs

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This paper represent a substantial extension and revision of Olper and Raimondi (2006). We benefited from discussions with participants at the EAAE 11th Congress in Copenhagen; the IATRC Meeting in Bonn and the IAAE Congress at the Gold Coast, Australia. This project was supported by the Università degli Studi di Milano (FIRST research grant: "Accordi commerciali internazionali, politica agricola comune e problematiche di accesso al mercato"). Part of this study was carried out when the first author was visiting professor at LICOS (Centre for Institutions and Economic Performance), University of Leuven (KUL). I wish to thank the LICOS staff, and especially Jo Swinnen, for their hospitality.

## 1. Introduction

Countries and governments are expressing growing concern with regard to reciprocity or asymmetry in market access caused by the differences in the structure of tariffs and the proliferation of standards, sanitary and technical barriers to trade.<sup>1</sup> The trade war between the EU and the US on ‘hormone beef’, GMOs and bananas, as well as that between the US and Japan on cars and apples, represent some illustrate examples of these issues. In theory, the use of a micro funded framework to measure trade policy restrictiveness (see Anderson and Neary, 2005) offers a sound approach to handle such problems. However, the approach is very demanding in that it requires high quality data on *direct* measures of policy barriers that, when existing, are poor and incomplete especially (but not only) with regard to non-tariff barriers to trade (see Bureau and Salvatici, 2004a; Anderson and van Wincoop, 2004).

Indeed, the evidence of *direct* protection measures often differs quite markedly, rendering the apparently simple question on ‘how high are policy barriers to trade?’ a very difficult answer when you are interested in a careful examination for most industries in most countries for several years. A look at the literature on EU agri-food protection is illustrative of this problem (see Bureau and Salvatici, 2004a). Direct estimates of the EU average agricultural tariff reveals a spread of estimates, ranging from the 40% of Messerlin (2001) to the 9.7% of Gallezot (2003), with intermediate value of 18% of Bureau and Salvatici (2004b). As suggested by the last authors, what is ‘puzzling’ here is that none of these figures are wrong, they simply result by different assumptions in calculating an aggregate tariff.

Given these problems, the literature now considers the possibility of using a complementary and *indirect* measure of market access, based on recent

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<sup>1</sup> This concern is well explained by the growing resource devoted by several countries in collecting and rationalizing ‘private’ information on market access issues. A review of the afford taken by the major trading blocs could be found in the last Report on the EU Market Access Database. See <http://madb.europa.eu/mkacddb2/indexPubli.htm>.

developments in the theory of gravity. Indeed this ‘workhorse’ tool has recently been developed in two important directions: the border effect approach initiated by McCallum (1995) and Wei (1996), and a more careful theoretical foundation due to Deardorff (1998), Head and Mayer (2000), Anderson and van Wincoop (2003) and many others.

Starting from these considerations, this paper uses the border effect estimated from a theory-consistent gravity model to assess the level and reciprocity in market access in food manufacturing trade among the United States, Canada, the European Union and Japan, the so-called QUAD countries. For a given region, the gravity model estimates the impacts of border effects (trade policy, preferences, information costs, etc.) on the relative volume of intra- versus international trade.

The estimation of market access through the border effect methodology offers two main advantages with respect to direct protection measures. First, as pointed out by Mayer and Zignago (2005), this approach accounts for the important and often neglected fact that internal demand is, for most products, largely met by domestic producers, not foreign ones. So an ideal protection index from the perspective of foreign producers needs a benchmark based on the best possible market access situation, i.e. that faced by national producers on the home market. This is exactly what the border effect approach reveals on comparing the relative volumes of intra- and international trade in two ‘identical’ countries. Secondly, the border effect methodology captures all the trade impediments related to the existence of national barriers, quite an advantage as most impediments are very hard to measure directly. For instance, consider the lack of reliable statistics concerning technical, sanitary and phytosanitary barriers that are so pervasive in the agri-food markets, and the inherent difficulty in estimating their trade effect. By using an overall picture based on an indirect estimation approach we overcome these problems.

Finally, the estimation of an overall market access index for sectors and countries characterized by a high degree of policy related barriers – like QUAD food markets – can offer new insights into literature that addresses the question of why national borders matter so much for international trade (see Hillberry, 1999; 2002; Evan, 2003; Anderson and van Wincoop, 2003; Chen, 2004; Fontagné

et al., 2005). Put differently, a contribution of this paper will also be to re-challenge earlier evidences concerning the main significance of the border effect, contrasting the trade policy view with other potential explanations such as ‘home bias’ in preference and spatial clustering.

From a methodological point of view, our analysis builds on recent papers of Fontagné et al. (2005) and Mayer and Zignago (2005). They applied the so-called log odd ratio specification of the Dixit-Stiglitz (1977) and Krugman (1980) ‘love of variety’ monopolistic competition trade model, first proposed by Head and Mayer (2000). This model yields a quite simple estimable prediction, derived directly from theory, for bilateral trade flows. The main advantage over other theory-driven gravity equations (e.g. Anderson and van Wincoop, 2003) is that this model allows an estimation of *reciprocal* bilateral border effects, one of the main objectives of our analysis.

The paper adds to the literature applying this approach to processed foods trade at a disaggregated level, an industry that is rarely covered by previous studies. This is an advantage, because analysis at aggregate level can mask, due to aggregation bias, important underlying phenomena linked to the inherent heterogeneity of the structure and the location behavior of the industries (see Hillberry, 2002). Differently, working at a more disaggregate level, should increase our ability to disentangle between competitive explanations on the determinants of border effect.

A second contribution of the analysis is our use of more comprehensive data with regard to policy trade barriers, especially non-tariff barriers (NTBs) to trade, in explaining the border effects. Specifically, advantage is taken of the *ad-valorem* equivalent of NTBs recently estimated by Kee et al. (2006) at HS 6-digit level. These new dataset offers two key advantages. First, by using the *ad-valorem* equivalent of NTBs, instead of frequency-type measures used in the previous literature (e.g. Fontagné et al. 2005), we overcame the problem of imposing a strong regularity condition in the estimation of the ‘trade’ effect of NTBs. This is because frequency-type measures assume that all changes in the index are equally important for trade flow, ignoring the fact that different NTBs can have different trade effects (see Anderson and van Wincoop 2004). Secondly, by using the *ad-valorem* equivalent of NTBs together with bilateral tariffs we

can directly compare their differentiated effect. This point could be important, as NTBs are trade policy components extraneous to actual WTO negotiations. Thus in order to understand the potential gain from trade policy reform, there is a critical need for a deeper comprehension of the differentiated effect of NTBs with respect to tariffs.

The paper is organized as follows. Section 2 summarizes the theoretical and empirical framework. Section 3 describes the data sources and the variables used in the empirical model. Section 4 is devoted to the presentation of our estimate on market access reciprocity, while in Section 5 we try to explain it. The final Section outlines our main conclusions.

## 2. Conceptual framework

Our gravity equation is based on the monopolistic competition trade model of Dixit and Stiglitz (1977) and Krugman (1980). Monopolistic competition is not the only model that can be used to derive a gravity-like equation<sup>2</sup>. However, it seems the most appropriate to consider processed food trade among the QUADs, given the high level of market concentration in food industries (McCorriston and Sheldon, 2007), the large dominance of developed country in food trade transactions, and the high and growing importance of intra-industry trade (Sheldon, 2005). In the derivation of the gravity model, we follow the structure proposed by Head and Mayer (2000), that offers the key advantage of estimating, in a rigorous framework, *reciprocity* in market access between two trading partners.

### 2.1 The model

Let  $x_{ij}$  be the volume of goods produced in country  $j$  and consumed in  $i$ , and  $\sigma$  the elasticity of substitution between home and foreign varieties. The representative consumer CES utility function in the importing countries  $i$  is given by

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<sup>2</sup> The first theoretical derivation of a gravity-like model is due to Anderson (1979). Deardorff (1998) derived gravity equations from the Heckscher-Ohlin model, Bergstrand (1989) from models with monopolistic competition, while that of Eaton and Kortum (2002) is from Ricardian models.

$$U_i = \left( \sum_j n_j (a_{ij} x_{ij})^{\frac{\sigma-1}{\sigma}} \right)^{\frac{\sigma}{\sigma-1}} \quad (1)$$

where  $n_j$  are (endogenously determined) varieties, identical to the number of firms;  $a_{ij}$  is the country  $i$  consumers' preferences for country  $j$  products.

Prices in the importing and exporting countries are introduced in the model using the notion of 'iceberg' trade costs  $\tau_{ij}$ , where  $\tau_{ii} = 1$  and  $\tau_{ij} \geq 1$ . This means that  $\tau_{ij}$  units of product must be shipped to country  $i$  in order for one unit to arrive. Thus, the consumers in  $i$  face a price  $p_{ij}$  for a product imported by  $j$  equal to the product of the price  $p_j$  and the corresponding trade costs,  $p_{ij} = p_j \tau_{ij}$ .

Let  $Y_i = \sum_k n_k x_{ik} p_{ik}$  be the country  $i$  consumers total expenditure on goods from all sources (varieties)  $k$ , including the domestic source. By maximizing the utility function (1) subject to the usual budget constraint, we get the country  $i$  total demand for country  $j$  products, namely the value of their bilateral imports

$$m_{ij} \equiv \sum_k x_{ij} p_{ij} = \frac{a_{ij}^{\sigma-1} n_j (p_j \tau_{ij})^{1-\sigma}}{\sum_k a_{ik}^{\sigma-1} n_k (p_k \tau_{ik})^{1-\sigma}} Y_i. \quad (2)$$

Equation (2) could be simplified by noting that in the Dixit-Stiglitz monopolistic competition model the firms set a price as if they face a constant price elasticity of demand, equal to the elasticity of substitution between varieties,  $\sigma$ . Thus the prices  $p_j$  are expressed as a constant markup over the marginal cost,  $p_j = c_j \sigma / (\sigma - 1)$ . It follows that the output per firm  $q$ , defined by the zero profit condition, also depends only on the elasticity of substitution and the marginal costs. Assuming these parameters to be the same across countries, results in proportionality between production value and the number of varieties  $n_j$ . Denoting  $v_j$  the value of production in  $j$ , we have  $v_j = q n_j p_j$ . This equation will be used to eliminate the unobserved  $n_j$  from the bilateral imports equation (2), using  $n_j p_j = v_j / q$ . Plugging this notation in (2) and taking logs, yields

$$\begin{aligned} \log m_{ij} = & \log Y_i + \log v_j - (\sigma - 1) \log \tau_{ij} - \sigma \log p_j + (\sigma - 1) \log a_{ij} - \\ & - \log \left\{ \sum_k a_{ik}^{\sigma-1} v_k (p_k \tau_{ik})^{1-\sigma} \right\} \end{aligned} \quad (3)$$



The first two terms of equation (3) capture the effects of country size, namely the industrial demand of the importer and the industrial supply of the exporter. The third term captures the effect of bilateral trade costs, while the fourth and fifth terms capture international differences in prices and preferences. The final term is the log summation of some highly non linear term related to variables of all countries, and come from the denominator of the CES price index. This term has been defined by Head and Mayer (2000) as the importer’s ‘inclusive value’ and by Lai and Trefler (2004) as ‘market potential’, because it depends on the full range of potential supplier  $k$  to a given importer.

The correct estimation of the bilateral import equation (3) needs to take account of the influence of this strongly non linear term, that clearly depends on parameters that are already in the equation to be estimated. However the importance of this term also depends on the focus of the empirical analysis. Indeed, if the objective of the researcher is to exploit the full general equilibrium response of trade flow to trade policies, as, for example, in Lai and Trefler (2004), then one needs to fully understand, empirically, its influence. On the other hand if the objective is less ambitious, one can follow the interesting manipulation first proposed by Head and Mayer (2000), working with the so-called log odds specification.<sup>3</sup>

To do that, we simply need to set  $j = i$  in equation (3) to allow an equation for country imports from itself,  $\log(m_{ii})$ . Then, the elimination of the non linear unobserved price index of the importing country, as well as its total expenditure, will be obtained by subtracting from equation (3) the specification of the country’s imports from itself, yielding

$$\log\left(\frac{m_{ij}}{m_{ii}}\right) = \log\left(\frac{v_j}{v_i}\right) - (\sigma - 1)\log\left(\frac{\tau_{ij}}{\tau_{ii}}\right) - \sigma\log\left(\frac{p_j}{p_i}\right) + (\sigma - 1)\log\left(\frac{a_{ij}}{a_{ii}}\right). \quad (4)$$

Equation (4) establishes a relation between the relative amounts consumers spend on foreign and domestic goods, and their relative price net of transport costs, and represents the theoretical counterpart of our empirical specification.

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<sup>3</sup> Another theoretically-consistent way of estimating equation (3) under symmetric assumption, is to follow Anderson and van Wincoop (2003) and Feenstra (2004), including fixed effects for source

## 2.2 Empirical specification

Before deriving an estimable equation, it is necessary to model both the trade cost and preference components of equation (4). Two elements of bilateral trade costs are considered: physical transport costs,  $d_{ij}$ , proportional to distances from  $i$  and  $j$ , and costs due to the presence of an international border,  $b_{ij}$ , such as tariff, non-tariff barriers, and other border related costs

$$\tau_{ij} = d_{ij}^{\rho} b_{ij} \quad (5)$$

where  $(b_{ij} - 1)$  is the tariff equivalent of all trade barriers associated with the border. Following Fontagnè et al. (2005), we assume that the structure of protection varies across all the partner pairs, and depends on the direction of trade flows

$$b_{ij} = (1 + t_{ij})(1 + NTB_{ij}) \equiv \exp(\eta EU_{ij} + \phi EUCAN_{ij} + \phi CANEU_{ij}). \quad (6)$$

For simplicity purposes, equation (6) considers only trade between European countries and Canada, where  $EU_{ij}$  is a dummy variable equal to 1 when  $i$  and  $j$  are both EU countries (for  $i \neq j$ );  $EUCAN_{ij}$  is a dummy variable equal to 1 when  $i$  ( $\neq j$ ) belongs to EU and  $j$  is Canada; and  $CANEU_{ij}$  is a dummy variable equal to 1 when  $i$  is Canada and  $j$  ( $\neq i$ ) belongs to the EU.

In the specification of consumer preferences,  $\alpha_{ij}$ , we simply assume that consumers prefer goods produced in a contiguous country ( $C_{ij}$ ) and that the sharing of cultural features ( $L_{ij}$ ) generates greater similarity in taste.  $L_{ij}$  and  $C_{ij}$  are two dummy variables that take a value 1 when country  $i$  and country  $j$  (for  $i \neq j$ ) speak a common language and/or share a common border (0 otherwise). Moreover, following Disdier and Mayer (2005) we also introduced variable proxies for common cultural traits and information related costs ( $I_{ij}$ ). This allows for a better control of preferences and other not-specified border components. The last variables are proxy by the bilateral flow of *books* and *newspapers* expressed as the value share available in the importing countries, also allowing control of the fact that when working with the EU countries the usual common language

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and destination countries. However, this set-up precludes the estimation of *reciprocity* in border

dummy does not offer a satisfactory proxy to capture differences in preferences and information related costs.<sup>4</sup>

Plugging in (4) the specifications of the trade costs (5)-(6) and taking into account consumer preference proxies, yields the following log odds form of the (relative) bilateral trade equation

$$\begin{aligned} \log\left(\frac{m_{ij}}{m_{ii}}\right) = & \log\left(\frac{v_j}{v_i}\right) - (\sigma - 1)\rho \log\left(\frac{d_{ij}}{d_{ii}}\right) + (\sigma - 1)\lambda L_{ij} + (\sigma - 1)\theta C_{ij} + (\sigma - 1)\xi I_{ij} - \\ & - \sigma \log\left(\frac{p_j}{p_i}\right) - (\sigma - 1)\eta EU_{ij} - (\sigma - 1)\phi EUCAN_{ij} - (\sigma - 1)\phi CANEU_{ij} + \varepsilon_{ij}. \end{aligned} \quad (7)$$

Taking the antilog of (each) estimated border dummy coefficients gives the reciprocal border effect of the respective trade flow combinations, namely the degree to which intra-country trade exceeds international trade after controlling for size, transport costs, preferences and relative prices. The Border effect dummies capture both the average level of protection of the importing country and any other not specified trade costs between  $i$  and  $j$ . For example, the coefficient on the  $EUCAN_{ij}$  dummy equal to  $(\sigma - 1)\phi$ , indicates the difficulty faced by Canadian exporters when selling its products to EU markets. Symmetrically, the coefficient on  $CANEU_{ij}$  equal to  $(\sigma - 1)\phi$  indicates the difficulty European exporters have in accessing the Canadian market. By comparing these estimated coefficients, we are able to examine potential asymmetries in market access, both at the overall country level and at the industry-specific level.

### 3. Data and measures

Our gravity model includes trade among 13 countries, 10 European Union<sup>5</sup> plus the other three QUADs, and 18 food industries. The database considers the imports of the QUAD countries for the period 1996-2001. The full data set presents a total of 14,125 observations and considers 33% of world food trade and 52% of QUAD country food imports from the world.

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effect, that represents the main objective of the present paper.

<sup>4</sup> The simple reason is that for the EU countries the language dummy displays very low cross-country variation.

<sup>5</sup> The observations for Austria, Greece, Ireland, Luxembourg and the Netherlands are not considered due to a large zero value in the production data.

We used the *UN Comtrade* database for bilateral trade at the HS-96 6-digit level, reported by the importer countries, then aggregated at the 4-digit ISIC industry classification. Output data at the same classification and aggregation level (ISIC rev. 3 code from 1511 to 1600) come from the *OECD Structural Statistics for Industry and Services database*, supplemented by other national sources in the case of missing values.

An early problem in the empirical implementation of equation (7) was that we needed intra-country trade data, but these figures were not available for our country sample. Thus, as in Wei (1996), Chen (2004) and others, we constructed such data on the basis of the assumption that what a country imports from itself is the difference between its total output and its total export to the rest of the world in each sector,  $m_{ii} = v_i - exp_i$  where  $exp_i \equiv \sum_{j \neq i} exp_{ij}$

Moreover, the implementation of the model also needs a measure of intra-national distances,  $d_{ii}$ . This introduces a not negligible difficulty in the model as any incorrect measure of intra-national distances translates into a proportional bias of the border effect estimate (see, Wei, 1996; Head and Mayer, 2002). To overcome this problem, we followed the approach implemented by Chen (2004) and Head and Mayer (2002), where distances between and within countries are computed using the same methodology. More specifically, both international and intra-national distances are based on bilateral distances between cities weighted by the share of the city in the overall country's population. Data on distances, together with the dummies on language and contiguity, are taken from *CEPII* (Centre d'Etude Prospectives et d'Informations Internationales).<sup>6</sup>

Relative prices are proxied by the price level of GDP expressed relative to the United States, taken from the *Penn World Tables v.6.1*. This choice was dictated by the low data availability of the industry-level prices and the underlying endogeneity issues.<sup>7</sup> Finally, as discussed in Section (2), we used two variables as proxy for information related costs and cultural proximity: bilateral flows of *newspapers* and of printed *books*, respectively. Both variables are calculated as bilateral imports relative to the importer production value. Trade data come

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<sup>6</sup> See, [www.cepii.fr/anglaisgraph/bdd/distances](http://www.cepii.fr/anglaisgraph/bdd/distances).

<sup>7</sup> See Fontagné et al. (2005) for an in-depth discussion of the *pro* and *cons* of this proxy.

from *UN Comtrade*, while the production data are from *OECD*, *UNIDO* and National statistics.

## 4. Results

### 4.1. Estimation of reciprocal market access and asymmetries

We start by estimating the bilateral magnitude of border effects, testing some different specifications based on the gravity equation (7). The results are reported in Table 1. Regression (1) reports our basic specification where we omit both books and newspapers, vector  $I_{ij}$  in equation (7), though these proxies are included in regression (2). Regression (3) adds two classical endowment measures to the specification, land per-capita and GDP per-capita, in order to control for differences in comparative advantage due to the resource based (land) nature of many food products. All the equations are estimated by ordinary least squares (OLS) pooled over 1996-2001, and across 18 food industries.<sup>8</sup>

The overall fit of the regressions are in line with the usual findings based on this gravity specification.<sup>9</sup> The coefficient on relative production, ranging from 0.84 to 0.75, is quite near the unitary value predicted by the model<sup>10</sup>. The trade elasticity of relative distance is, as expected, negative and strongly significant, with a value close to  $-1$  in all the specifications. In equation (1) the relative price coefficient is positive, though insignificant. However, it turns out to be negative and significant in regressions (2) and (3) when we control for ‘preference’ and resource endowment. Finally, the coefficients on contiguity and language are, as expected, positive and significant, suggesting that two countries sharing a

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<sup>8</sup> We also experimented a Heckman two-stage procedure to address selection bias concerns, due to the zero value of some bilateral trade flow combinations. Specifically, the dataset has 7.5% zero trade flows. However, the Mills ratio on the second stage is never significant, suggesting that selection bias concerns are not a problem here.

<sup>9</sup> Specifically, the relative nature of our specification gives lower explanatory power with respect to traditional or fixed effect gravity specifications. This is not surprising as in our case the variables are computed as differences with respect to internal flow used as the reference. Thus, our specification is comparable to a first-difference panel model estimation that, notoriously, increases the variance to be explained compared to the estimations in levels (see Combes et al., 2005 on this point).

<sup>10</sup> However, these coefficients are always significantly different from 1 (Wald test). We also ran a regression with this constraint included, by shifting relative production to the left side of the equation (see Head and Mayer, 2000). However, as this specification does not, to any degree, alter the results and conclusions of the paper, we saved space by not reporting these additional results.

common language and/or a common border tend to trade more with each other than otherwise.

All the border dummy coefficients are negative and strongly significant. This means that, after controlling for the bilateral determinants of trade flows suggested by theory, intra-country trade is significantly larger than cross-border trade. In regression (1) the coefficients for intra-EU trade equals  $-4.19$ , meaning that intra-country trade is, on average, 66 times greater [ $=\exp(4.19)$ ] than crossing a national border between EU countries. Strictly speaking, a comparable estimate for processed food trade does not exist. However, Furtan and van Melle (2004) detected even larger border effects in agri-food trade among NAFTA countries. Moreover, our results are in line with those of Head and Mayer (2000) and Chen (2004) who show higher intra-EU border coefficients for most ‘ingestible products’.

Contrary to expectations, the border effect for intra-EU trade is not the lowest among QUAD countries. Japan’s market presents a higher access level for imports from all the countries considered here, especially from the United States. The results, albeit quite surprising, are similar to previous findings conducted at the manufacturing level. Fontagné et al. (2005) consider that this ‘spectacular’ result might be driven by an overestimate of the US-Japan distance with respect to intra-EU distances. However, another explanation could be related to comparative advantage differences, due to the resource (land) based nature of many food products. We will return to this point later.

Equation (2) adds to the specification our proxy for ‘preference’ and information related costs, namely bilateral trade in books and newspapers. Both variables are positive and strongly significant and, interestingly, they induce a relevant reduction in all bilateral border coefficients, as well as in other trade cost coefficients (i.e. language, contiguity and distance). This is because, controlling for preferences and information related costs, the border coefficients are now ‘cleansed’ of these sources of market access differences, as previously shown by Olper and Raimondi (2006).

Finally, equation (3) adds two classical endowment measures, land per-capita and GDP per-capita, to control for cross-country differences in comparative advantage. Indeed, it is quite natural to think that bilateral food trade is

positively affected by the relative endowment in land, given the resource based (land) nature of many food products, and by the overall differences in capital per worker. The two variables have the expected positive sign and are highly significant and, interestingly, they induce a significant change in the estimated border coefficients that appears in line with the factor-proportion-comparative-advantage hypothesis. Indeed, in each bilateral combination we see a reduction (increase) in the border coefficients of the relative land *abundant* (scarce) countries that is roughly proportional to the country's endowment differences in land.

With this modification of the base model, the strong border effect asymmetries shown in regression (1) are partially recomposed with common perception. For example, the figures suggest that there exists a higher border protection in the EU with respect to US producers, rather than the reverse. More specifically, the border effect of US export to EU, equal to 137 [=exp(4.93)], strongly contrasts with the reciprocal border effect of EU export to US of only 15.7 [=exp(2.75)]. A similar story can be seen for the border effect of US export to Canada, equal to 181 [=exp(5.20)] with respect to the reciprocal one equal to only 22.7 [=exp(3.12)]. More symmetry in the reciprocal market access is detected for trade flows between the EU and Canada and US and Japan. In the first case the data show a border effect of 94 for Canadian exports to the EU market and of 58 for the flow in the opposite direction. In the second case, the border effect of US exports to Japan is only 11.8, while the reverse equals 25.7.<sup>11</sup> Moreover, controlling for 'preferences' and comparative advantage, the border effect for intra-EU food trade is about 19.3, a figure not so far from the findings of Fontagné et al. (2005), who found a border effect of about 12.8 for all manufacturing industry.

Finally, note that Japan's border effect incongruence, albeit reduced, is still present, and appears quite strong, especially in trade combinations involving Canada and the EU as trading partners. As Japan is the world's leading food importer – due to its strong natural resource (land) constraints – a probable explanation of the above results also lies in the fact that our monopolistic competition model is not well suited to capture this effect.

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<sup>11</sup> All the estimated reciprocal border effect differences between the country pairs (asymmetry) are statistically significant at 1% level (Wald test).

#### 4.2. Industry level market access

The border effect also differs across industry. Table 2 gives the results of industry-specific gravity equations with a specification that omits our proxy for information related costs and ‘preferences’, to estimate ‘overall’ differences in industry border effects. The results are based on OLS or Heckman two-stage procedure to address selection bias concerns. Specifically, when in the Heckman’s second stage regression the Mills ratio was significant, we rely on this estimated results, differently we use results from simple OLS regressions<sup>12</sup>.

One advantage of these industry regressions is that the trade elasticity to distance is now also industry-specific, thus it captures the across sector differences in physical transport costs, as well as in the elasticity of substitution. Due to space constraints the table reports only the estimated coefficients of distance and border dummy, with their respective *t*-value, and the implied value of border effect, sorting the food sectors in decreasing order of magnitude<sup>13</sup>.

In these equations the border coefficient is the constant. The highest border effect value, equal to 696, is found for the Manufacture of Sugar, followed by the Manufacture of Prepared Animal Feed, Tobacco and Manufacture of Malt Liquors. On the contrary, the lowest border effects are estimated for the Distilling Industry, the Manufacture of Starch products and Fish, where the border effects range from 4.3 to 9, respectively. These results are in line with those estimated by Head and Mayer (2000) and Chen (2004) on intra-EU trade, both for order and magnitude of border effects, suggesting that different countries display similar patterns in food industry border effects.

As it is evident from Table 2, the distance elasticity varies significantly across industry, ranging from  $-2.20$  for Manufacture of Grain Mill Products to  $-0.96$  for Manufacture of Prepared Animal Feed, with an across industry average value of

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<sup>12</sup> The Heckman two-stage procedure was implemented as follows. We first estimate a probit model regressing, on an indicator of positive trade, all the explanatory variables of the model plus levels of the exporter’s production, price, and distance (see Head and Mayer, 2000). Then, from the probit coefficients, we calculate the Mills ratio adding them to the original (OLS) specification. The probit estimated coefficients (not reported) are quite similar in signs and significance to the OLS. Finally, as a general rule the Heckman procedure tend to increase somewhat the estimate border coefficients.

<sup>13</sup> The (not reported) coefficients on relative production, contiguity, language and land *per-capita*, are systematically significant and display the correct sign; on the other hand relative prices and GDP *per-capita* in a few regressions are insignificant and/or of the wrong sign.



-1.71.<sup>14</sup> Most interesting is the fact that distance elasticity shows a strong negative correlation with industry border coefficients ( $r = -0.66$ ). This is a quite counterintuitive result. In fact, from equation (7) we know that the coefficients of distances are given by the elasticity of substitution times the elasticity of trade costs with respect to distance. Thus, if the substitution elasticity represents a large component of the border effect explanation, we can expect distance and border coefficients to be positively correlated (see Chen, 2004: 106). However, exactly the opposite was found. So, how to explain this negative correlation? An appealing explanation is one of political economy motives for trade protection. Indeed, countries give higher protection to industries where physical transport costs are less severe, and tend to give less protection to sectors where, due to high transport costs, import competition is less important.

Estimations at the industry level allow us to evaluate the degree of reciprocity in QUADs bilateral market access. To this end, we split each industry border dummy into 13 border dummies, one for each bilateral relationships between the United States, Canada, the EU and Japan. Figures 1 to 6 show the results, borrowing the representation approach of Fontagné et al. (2005). Each figure displays the reciprocal (absolute) border coefficients of a particular trade combination, sorted by the increasing value of market access in one direction. For example, in figure 1 the square dot represents the industry border effects faced by EU exporters on US markets, sorted by their increasing magnitude, while the triangular dot represent the reciprocity border effect faced by US exporters on the EU markets. The vertical positions of the squares and triangles, show the level of asymmetry in market access for each industry.

From a preliminary inspection of the six figures, two main patterns emerge: a positive correlation between reciprocal market access of different industries for several combinations; a strong asymmetry in food industry border effects across country pairs. Starting from the first point, the positive correlation in market access is very high for trade between Canada and the US ( $r = 0.77$ ), and between

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<sup>14</sup> This average elasticity is significantly higher in absolute value than the distance elasticity shows in Table 1. This is due to the different specifications used here, where we omit the country border dummies to estimate (across-country) average industry-specific border effects. In the industry regressions that also include the bilateral country dummies (see below), the across industry average distance elasticity is equal to -1.17. The correlation between the two series of distance

the EU and Japan (0.64). It assumes an average value for trade between the EU and the US (0.52) and the EU and Canada (0.32), but is virtually zero for trade between Japan and both the US and Canada. A similar, but less important, positive correlation was detected by Fontagné et al. (2005) at the manufacturing level. They interpret this pattern in terms of endogenous protection, namely industries in different countries have the same pattern of sensitivity, probably because they display similar industry characteristics. Moreover, because a component of our market access proxy is ‘home bias’ in preferences, that is notoriously quite strong in processed foods (see Lopez et al., 2006), another explanation for this positive correlation could be the similar degree of ‘home bias’ in preferences across industries in different countries.

Let us move on to the results for reciprocity in market access. From the figures it appears quite evident that in all the combinations we detect significant asymmetry in border effects. That is to say, several countries tend to impose higher border costs on the other partners for the majority of processed foods. For example, this is the case for trade between the EUs *vs.* the US, and for Canada *vs.* the US (see Figures 1 and 2). In the first case, the only sectors where the EU is more open than the US is the Feed industry; in the second case Canada is more open than the US only in the Wine industry. There seems to be a little more symmetry in market access between the EU and Canada, with the EU more open than Canada in the Distilling and Blending of Spirits, Manufacture of Other Food, Malt/Liquors and Sugar, and with very similar market access in the Feed industry, Manufacture of Meat and Soft Drinks (see Figure 3). On the other hand, the only country pair combination that displays some form of symmetry in market access is that between the US and Japan (Figure 4), where we observe a relatively ‘comparable’ number of positive and negative differences in border effects. Note however that, all the country pair combinations involving Japan, suffer the ‘puzzling’ result discussed above, that sees Japan often more open to food imports than the reverse, and this is especially true for combinations involving the EU (Figure 5).

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elasticity is close to 1, showing that the across industry ranking in physical transport costs reported in Table 2 is informative.

At the level of specific industries, it appears clear that the sectors of Distilling and Blending of Spirits on the one hand, and that of Manufacture of Sugar on the other, represent two systematic outliers in the majority of country pair combinations. Sugar is characterized by a very large border effect, while Spirits normally display a border effect not significantly different from zero, and in the case of Canada's exports to Japan, there is a positive and significant border effect (Figure 6), suggesting a preference for trading with other countries rather than with itself. The data show only one other case of a significant positive border effect, that for US imports of Fish from Japan. Finally, we detected 7 other border effects not significantly different from zero, all in combinations involving Japan with the US or Canada.

## **5. What determines these big asymmetries?**

In the previous section we documented a significant variation in border effects both among countries and across industries. The objective of this section is to collect proxies for the main determinants of border effects with the aim of re-challenging previous findings.

Theoretically, border effect is equal to the product of two things (see Section 2.2): the degree of substitutability between goods and the tariff equivalent of border related costs. Both terms are crucial to understand border effects. In sectors with high substitutability between domestic and foreign goods, even low level of border barriers can induce a significant reduction in market access. Differently, for a given value of the elasticity of substitution, each country or industry characteristic that affects the border barriers represents a good candidate to explain the border (Chen, 2004). Here, we focus our attention on four key potential explanations: trade policy (tariffs and NTBs), spatial clustering, the degree of product differentiation (as proxy for elasticity of substitution), and 'home bias' in preferences.

Clearly, the more intuitive explanation is that border effect largely arises because governments use trade policy to discriminate between domestic and foreign goods. However, previous evidence does not univocally support this interpretation. For example, Hillberry (1999), for trade between the US and Canada, and Head and Mayer (2000) and Chen (2004) for intra-EU trade, did not

find any significant role for tariff and non-tariff barriers, respectively. At the same time, though some papers do find a role for policy (see Evans, 2003; Fontagné et al., 2005), the component of the border attributable to trade policy is quite tiny.<sup>15</sup>

To test the trade–policy–border–effects relationship we make use of industry *ad valorem* bilateral tariffs computed starting from the Market Access Map (MacMap) tariffs at HS 6-digit (see Boüet et al., 2004), then aggregated using import weights to match our ISIC 4-digit classification. For non-tariff barriers (NTBs) we use the *ad-valorem* equivalents (AVE) recently estimated by Kee et al. (2006) at the HS 6-digit level,<sup>16</sup> aggregated at the ISIC 4-digit, using import weights. It is important to note that, due to data problems, all the previous applications instead of the AVE of NTBs, have normally used or simple dummy variable (Head and Mayer, 2000; Chen, 2004), or frequency-type index (Haveman et al., 2003; Fontagné et al., 2005; Mayer and Zignago, 2005) to capture the trade effect of NTBs. However, this procedure is not without problems as it imposes strong regularity conditions. Indeed, it assumed that all changes in the coverage ratio are equally important for trade flow, ignoring the fact that different NTBs can have different trade effects (see Anderson and van Wincoop 2004). Moreover, an interesting advantage in using the AVE of NTBs lies in the direct comparability with the estimated coefficient of tariffs.

Let us now move on to other potential explanations, considering spatial clustering first. Several authors suggest that border effect could arise endogenously because, in order to avoid trade costs, firms producing intermediate and final goods tend to co-agglomerate within political boundaries (see Wolf, 2000; Hillberry and Hummels, 2002). The argument is that firms not tied to a specific location tend to locate to minimize trade costs, thus reducing

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<sup>15</sup> One explanation for this weak evidence could, however, be data problems: actual direct protection measures only partially capture the whole story. For example, we lack good data on several non-tariff barriers to trade, such as the myriad of technical and sanitary barriers imposed by government regulations. At the same time tariff data, notoriously less problematic than NTBs, are not without problems, especially when we look at their time dimension.

<sup>16</sup> To obtain the AVE of NTBs, these authors first estimate the impact of NTBs on imports at the tariff line level using Leamer’s comparative advantage approach. Then they transform the quantity impact into price equivalents, using a careful estimate of the import demand elasticity, at the HS 6-digit level. Note that, these data are not without problems (see Ferrantino, 2006). For example, they ignore the potential positive effect on trade flows of some NTBs, due to the increase of information diffusion and network externality. However, to date, they represent the most accurate AVE of NTB estimates, with a large country/industry coverage.

international trade and increasing (endogenously) the border effects. The applicability of this reasoning to food industry, however, is not without problems. Indeed, food sectors are often linked to a particular location not only with the intention of saving in transport costs, but also as a consequence of natural resource advantages (think of the wine industry) or because they want to satisfy customers' face-to-face requirements.

In order to test the possible 'endogeneity' of the border effects, we make use of an industry, distance-based, 'spatial concentration' index, computed by Busch and Reinhardt (1999)<sup>17</sup> for US data at SIC 4-digit level, then converted to ISIC 4-digit to match our industry classification. Unfortunately, the intensive data requirements to construct this index, prevented our use of country specific indices. However, previous authors who took a similar approach (see Chen, 2004) found robust results by adapting the US geographic concentration index to EU countries. The index is increasing in 'spatial concentration' with an average value for our industry sample of 0.43, a higher value for Manufacture of Tobacco (0.64) and a lower one for Processing and Preserving of Fish (0.22).

As discussed above, a key component of the border is the degree of product differentiation. Higher values tend to reduce the substitution possibility between domestic and imported goods. Thus trade will be higher, inducing a reduction in border effects. The degree of product differentiation is often proxy through the Rauch (1999) categorization of industry in 'homogenous', 'reference priced' and 'differentiated products' (see Evans, 2003; Chen, 2004).<sup>18</sup> However, processed foods fall in the same Rauch category (homogenous goods), preventing our discrimination in terms of the degree of product differentiation. To circumvent this problem, we used an intuitive, admittedly imperfect, proxy to measure the degree of product differentiation: the log number of HS 6-digit category in each ISIC 4-digit.<sup>19</sup> This number shows high variability in processed foods, ranging from 2 HS line for ISIC-1533 (Manufacture of Prepared Animal

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<sup>17</sup> The spatial concentration index is measured as a decreasing function of the distance from each employee to the national 'centroid' or mid point for a given industry, and is based on 1987 US data. See Busch and Reinhardt (1999) for data source and computation details.

<sup>18</sup> Strictly speaking, in Chen (2004) the Rauch index was intended to capture information related costs, on the grounds that 'search costs' tend to increase for differentiated products. Differently, Evans (2003) interprets the Rauch index as a proxy of the substitution elasticity across varieties.

<sup>19</sup> More specifically this index represents the 'potential' degree of product differentiation for each ISIC 4-digit, because no longer are all HS-6 lines always fulfilled in each country pairs.

Feed) to 75 HS line for ISIC-1513 (Processing and Preserving of Fruit), with an average lines number of 22.5.

Finally, to capture ‘home bias’ in preferences, we take advantage of the Broad Economic Categories (BEC), to classify products in intermediate *versus* consumption goods. Basically, using the BEC concordance with HS 6-digit, we classified our ISIC 4-digit industries in terms of the value share in each sector for a given importer who deals in consumption goods. The index ranges from 0 to 1, with an average value of 0.78. The underlying hypothesis is that preferences will be more distorted towards domestic products when consumers, rather than firms, decide the origin of the goods consumed (Fontagné et al., 2005).<sup>20</sup>

### 5.1. Estimation and Results

To estimate the impact of the various factors discussed above we follow the approach used by Head and Ries (2001), Head and Mayer (2000) and Chen (2004), who regress the country- and industry-specific border coefficients on their determinants. Despite not being ideal from an empirical point of view, due to the use of an econometric estimate as dependent variable, it allows a direct and simple comparison of the explanatory power of each supposed determinant of the border effects.<sup>21</sup>

In what follows, we estimate variations of the following basic equation derived from the model specification discussed in Section 2

$$b_{ik} = (\sigma - 1)\log(1 + t_{ik}) + (\sigma - 1)\log(1 + NTB_{ik}) + (\sigma - 1)\beta_{ik} + \varepsilon_{ik}$$

where  $b_{ik}$  are the 234 (13 x 18) estimated country- industry-specific border coefficients based on equation (7);  $\beta_{ik}$  represents a vector of all the determinants

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<sup>20</sup> A limitation of this index is that the level of tariff protection also increases on passing from primary, intermediate to final consumer goods in food industry, due to tariff escalation.

<sup>21</sup> Another approach is to directly insert each determinant of the border into the gravity equation (see Fontagné et al., 2005; Olper and Raimondi, 2006), and measure the resulting reduction in the estimated border coefficients. A problem with this approach is that we are comparing results from different estimations. In fact, the estimations do not include the same variables, and estimated coefficients of several explanatory variables (e.g., distance, contiguity and language) differ between the regressions. Therefore, part of the variation observed in the constants of the equation, which represent the border effects in our model, is just due to variations in the estimated coefficients on all variables between regressions.

of the border different from trade policies, comprise a common intercept and (eventually) country and industry fixed effects; finally  $\varepsilon_{ik}$  is the error term.

The estimation of the 234 border effects is based on one year's (2000-2001) OLS regression, as the key determinants of our interests, tariffs and AVE of NTBs refer to this period.<sup>22</sup> Finally, following Head and Mayer (2000) and Chen (2004) to reduce the heteroskedasticity problems of using a dependent variable estimated with error, we use weighted-least-squares, with weights equal to the inverse of the standard errors of the estimated border coefficients. Finally, to simplify the interpretation of the estimated coefficients, we have multiplied the dependent variable by  $-1$ .

The regression results are reported in Table 3. In line with our *a-priori* expectation, the border effects display a positive and significant association with trade policies, suggesting that both tariffs (Column (1)) and NTBs (Column (2)) are important components of the border<sup>23</sup>. It is interesting to note that NTBs display both a higher estimated coefficient and explanatory power than tariffs. Thus, for our country and industry sample, the border effects appear particularly affected by non-tariff barriers to trade. The degree of 'spatial concentration' displays a significant and positive affect on the border effects. This means that the more 'spatial concentrated' a sector is, more it display an higher, and not lower, border effect. A result contrary to Chen (2004) who shows that, in the EU, firms that are not geographically concentrated display larger border effects. Thus our findings do not support the argument of Wolf (2000), Hillberry (2002), and Hillberry and Hummels (2002) on the endogeneity of the border effects, suggesting that this kind of interpretation tends to be industry-specific. Moreover, a competing explanation with this argument, not mentioned in the previous literature on border effect, is that because 'spatial concentration' represents a (positive) political economy determinant of protection (see Busch

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<sup>22</sup> Clearly, it would be better to estimate these border coefficients through specific-industry regressions. However, when working on a single year certain regressions have very few observations, inducing volatility in estimated border effects.

<sup>23</sup> The coefficients on tariffs implies an average elasticity of substitution between varieties equal to 3.30. This value is somewhat higher than the recent estimate by Lai and Trefler (2004), who find an elasticity of substitution of 2.50 for aggregate food industry (3-digit ISIC). Thus, because we are working at a 4-digit level, our result is consistent with the notion that the elasticity of substitution tends to increase with the level of desegregation (see Anderson and van Wincoop, 2004).

and Reinhardt, 1999), the significant positive coefficients shown in Table 3 could simply capture this kind of relationship.<sup>24</sup>

Our proxy for the degree of goods differentiation displays the correct, and significant, negative effect (Column (4)). Thus, the more differentiated the goods, the less sensitive consumer is to a given level of trade costs, and the lower the border effects, a finding consistent with both Evans (2003) and Fontagné et al. (2005). Finally, Column (5) adds our proxy for ‘home bias’ in preferences, the share of final goods in each industry for the importer countries. It shows positive and significant coefficients, meaning that for intermediate goods the border effect level tends to be lower: thus, ‘home bias’ in preferences represents a part of the story. However, note that the explanatory power of our proxies for both the elasticity of substitution and ‘home bias’ are significantly lower than that of policy variables, suggesting that in the food industry trade policies matter.

Next, adding together all the explanatory variables (Column (6)), it can be seen that the coefficients on tariffs shrink substantially and become insignificant. This effect is due especially to the simultaneous introduction of NTBs and ‘spatial concentration’ index, that are in fact both positively correlated with tariffs. Finally, in Column (7) we add to the regression the industry fixed effects at 3-digit level to check for robustness. As it is clear by the inspection of the regression coefficients, our results and conclusions are substantial unaffected.

Overall, our results suggest that, in the food industry, trade barriers, especially in the form of NTBs, are important components of the reason why national borders affect trade flows. However, the results also highlight that elements not directly linked to trade costs, such as home bias in preferences and the elasticity of substitution, play a role in explain the magnitude of the border. Thus, the increase in market integration due to regional and WTO trade agreements should, through the removal of trade barriers, contribute to increased trade flow in the food industry, reducing border effects. An indirect implication arisen from the emerging importance of NTBs in explaining the borders. Indeed, because NTBs are trade policy components extraneous to actual

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<sup>24</sup> Introducing the ‘spatial concentration’ index together with tariffs induces a not marginal reduction of the coefficient and significance on tariffs, giving credence to this interpretation.



WTO negotiations, our figures allow a deeper integration strategy based especially on bilateral trade agreements, where elements such as 'mutual recognition' and harmonization issues play a central role.

## **6. Conclusions**

The present paper analyzes the level and asymmetry in food industry reciprocal market access of the so-called QUAD countries – the US, Canada the EU and Japan. These countries represent the main trading blocks that account for more than one third of world trade in processed foods. Using the border effects approach estimated from a structural gravity model, we measure the friction each country faces when exporting to another QUAD countries. In a second step we challenge the main explanation of border effects.

At the aggregate level we detected strong asymmetry in market access: indeed, US exports to the EU appear much more affected by the border than the reverse, and the same can be said for US exports to Canada. Quite surprisingly, Japan seems to be more open to processed food imports from QUAD countries than the reverse, and this holds true even if we control for differences in comparative advantage.

The industry-level analysis gives details on market access at the 4-digit ISIC level, confirming the presence of reciprocal asymmetric border effects and showing that countries tend to display quite similar patterns in across-industry variation in border effects.

Exploring the possible underlying reasons in border effects we show that trade policies, especially in the form of non-tariff barriers to trade, seem to matter in food industry, and this partially contrasts with previous findings. However, other underlying explanations of the border, especially 'home bias' in preferences and the magnitude of substitution elasticity, also play a role. These findings are broadly consistent with the most recent developments in border effect literature (e.g. Evans, 2003; Fontagné et al. 2005; Mayer and Zignago, 2005). On the other hand, we did not find any evidence that the border effect is 'endogenous', due to firm's location decisions in order to reduce trade costs, as suggested, among others, by Hillberry (2002) and Chen (2004). In the food industry, differently, the 'endogeneity' of border effect seems largely be due to political economy motives,

namely government-set trade policies to contrast import competition from sectors where, apart from trade policy, trade costs tend to be low.

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Table 1: Levels of Border Effects in Processed Foods

Dependent variable	Log relative import ( $m_{ij}/m_{ii}$ )					
	Eq. (1)		Eq. (2)		Eq. (3)	
	Coeff.	<i>t-value</i>	Coeff.	<i>t-value</i>	Coeff.	<i>t-value</i>
<i>Log rel. production</i>	0.84	75.08	0.75	62.86	0.79	62.71
<i>Log rel. distance</i>	-1.10	29.99	-1.01	28.21	-1.11	27.76
<i>Log rel. prices</i>	0.07	0.71	-0.22	2.29	-1.42	8.18
<i>Common Language</i>	0.95	11.94	0.40	5.08	0.42	5.31
<i>Contiguity</i>	1.11	21.50	0.84	16.83	0.76	14.87
<i>Log rel. books</i>			0.12	8.67	0.12	8.73
<i>Log rel. newspapers</i>			0.08	8.38	0.07	7.44
<i>Log rel. gdp-pc</i>					0.79	9.05
<i>Log rel. land-pc</i>					0.35	10.84
<i>EU to EU</i>	-4.19	63.37	-3.09	32.88	-2.96	30.40
<i>CAN to EU</i>	-5.47	36.15	-3.96	23.63	-4.54	26.47
<i>EU to CAN</i>	-6.52	62.51	-5.02	35.05	-4.06	23.72
<i>CAN to USA</i>	-4.54	30.53	-3.13	18.51	-3.12	18.45
<i>USA to CAN</i>	-6.51	47.50	-5.36	34.80	-5.20	33.29
<i>CAN to JPN</i>	-2.07	8.33	-1.00	3.86	-2.15	7.92
<i>JPN to CAN</i>	-8.02	33.92	-6.67	26.34	-5.08	17.33
<i>EU to JPN</i>	-3.44	24.08	-2.45	15.72	-2.76	17.79
<i>JPN to EU</i>	-7.73	47.84	-6.27	35.43	-5.38	25.56
<i>USA to EU</i>	-5.37	37.68	-4.45	29.28	-4.93	31.56
<i>EU to USA</i>	-4.94	49.40	-3.63	28.03	-2.75	17.87
<i>USA to JPN</i>	-1.65	8.57	-1.42	7.25	-2.47	11.84
<i>JPN to USA</i>	-5.90	27.23	-4.76	20.56	-3.25	11.95
N	14,125		14,125		14,125	
Adj R <sup>2</sup>	0.517		0.533		0.538	

Notes: Pooled OLS regressions; Each equation also includes a set of year fixed effects; *t*-values are corrected for Heteroskedasticity.

Table 2: Industry-specific border effects and distances elasticity

	ISIC	Distance		Border coefficient		Border effect	adj-R2	N.	Method
		Coeff.	t-value	Coeff.	t-value				
Manufacture of sugar	1542	-1.39	-6.38	-6.55	-12.04	696.5	0.55	429	OLS
Manufacture of prepared animal feed	1533	-0.96	-11.94	-5.12	-29.71	167.3	0.62	853	Heckman
Manufacture of tobacco products	1600	-1.53	-9.29	-4.88	-16.14	131.6	0.46	617	Heckman
Manufacture of malt liquors and malt	1553	-1.36	-12.88	-4.71	-20.27	111.1	0.51	738	Heckman
Manufacture of soft drinks and mineral water	1554	-1.71	-18.66	-4.61	-23.95	100.5	0.55	856	Heckman
Manufacture of vegetable and animal oils	1514	-1.55	-14.37	-3.52	-15.29	33.7	0.48	887	OLS
Manufacture of bakery products	1541	-1.77	-23.40	-3.43	-23.83	30.9	0.61	899	Heckman
Manufacture of macaroni, noodles, couscous...	1544	-1.84	-17.54	-3.23	-13.50	25.4	0.60	770	OLS
Manufacture of other food products n.e.c.	1549	-1.54	-26.43	-3.11	-23.22	22.4	0.64	935	OLS
Production, processing and preserving of meat	1511	-2.04	-18.94	-3.05	-14.77	21.1	0.59	879	Heckman
Manufacture of cocoa, chocolate and sugar conf.	1543	-2.06	-22.88	-2.91	-10.61	18.4	0.69	905	Heckman
Processing and preserving of fruit	1513	-1.63	-24.52	-2.71	-18.68	15.0	0.68	929	OLS
Manufacture of dairy products	1520	-2.17	-16.95	-2.51	-10.47	12.3	0.56	876	Heckman
Manufacture of grain mill products	1531	-2.20	-20.16	-2.49	-11.79	12.1	0.65	870	Heckman
Manufacture of wines	1552	-1.65	-12.73	-2.46	-8.05	11.7	0.76	507	OLS
Processing and preserving of fish	1512	-1.52	-19.33	-2.20	-13.06	9.0	0.60	780	Heckman
Manufacture of starches and starch products	1532	-1.82	-18.41	-2.16	-10.33	8.7	0.55	879	OLS
Distilling, rectifying and blending of spirits	1551	-2.10	-15.48	-1.47	-5.29	4.3	0.69	516	Heckman

Notes: OLS or Heckman second stage regression results; Each equation also includes a set of year fixed effects; Border effects are given by the antilog of the border coefficients; Coefficients (not reported) on relative production, distance, land per-capita are systematically significant and display the correct signs; otherwise relative prices and GDP per-capita in few regressions are insignificant and/or with wrong sign; *t*-values are corrected for Heteroskedasticity.

Table 3: Determinants of the border effect

	Dependent variable ( $b_{ik}$ )						
	Eq. (1)	Eq. (2)	Eq. (3)	Eq. (4)	Eq. (5)	Eq. (6)	Eq. (7)
Intercept	3.68 (18.31)	3.54 (17.51)	3.36 (6.96)	3.96 (8.19)	3.40 (12.31)	5.62 (14.31)	6.19 (7.76)
Log ( $tariff_{ik}$ )	2.30 (3.97)					0.36 (0.66)	0.42 (0.69)
Log ( $NTB_{ik}$ )		3.05 (5.13)				2.94 (5.29)	2.84 (4.85)
Log ( $conc_k$ )			2.08 (5.16)			2.46 (6.10)	2.34 (4.59)
Log ( $HS\ num_k$ )				-0.37 (4.15)		-0.40 (5.10)	-0.49 (3.48)
Share of consumer goods $_{ik}$					0.57 (2.21)	1.30 (4.79)	1.05 (2.42)
Country fixed effects	yes	yes	yes	yes	yes	yes	yes
Industry (3-digit) fixed effects	no	no	no	no	no	no	yes
N	233	233	233	233	233	233	233
Adj $R^2$	0.132	0.187	0.137	0.086	0.057	0.392	0.388

Notes: Weighted last-square estimations, see text for details; *t*-values are given in parentheses; the reported adj  $R^2$  refer to un-weighted statistics; moving from the left to the right the weighted adj  $R^2$  are, 0.397, 0.422, 0.398, 0.401 0.365, 0.572 and 0.568, respectively.

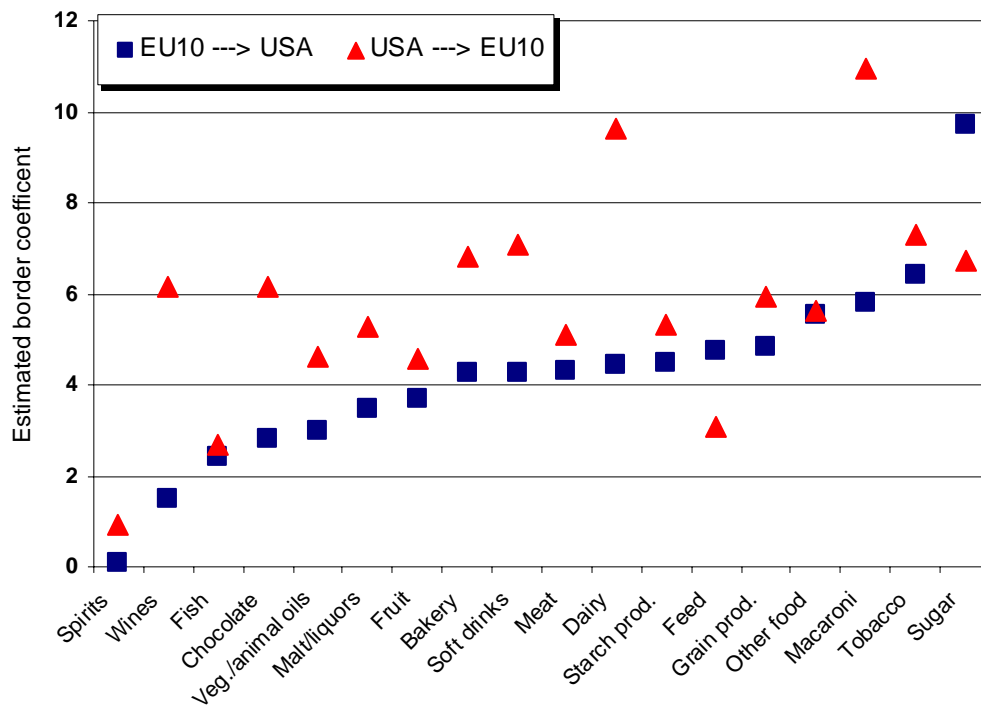


Figure 1: Market access asymmetry across industry between the US and the EU10

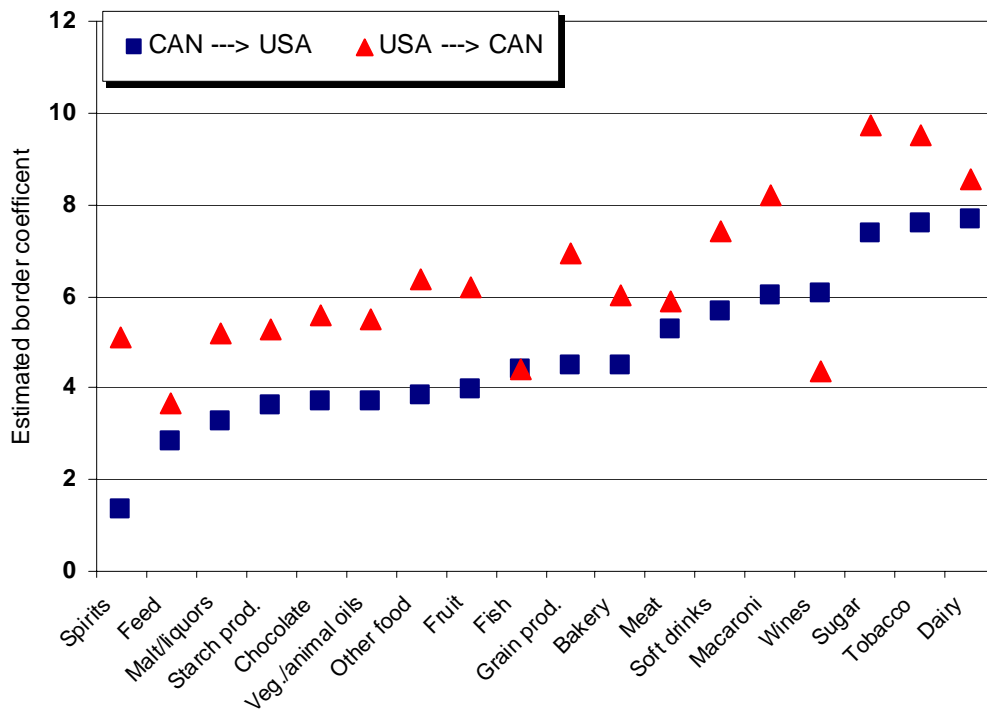


Figure 2: Market access asymmetry across industry between Canada and the US

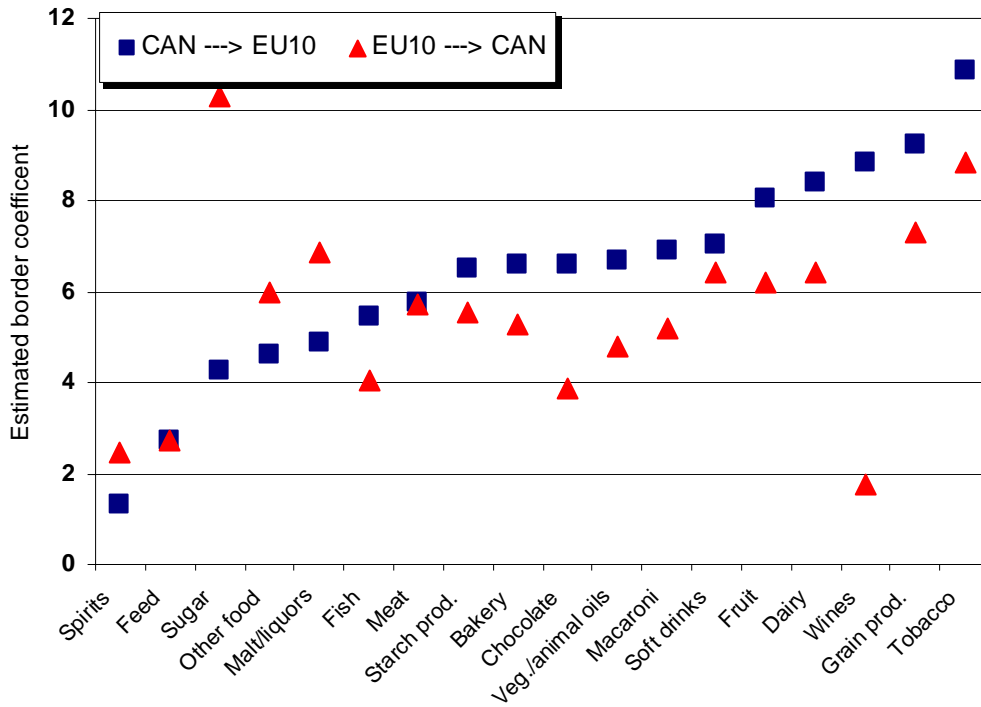


Figure 3: Market access asymmetry across industry between Canada and the EU10

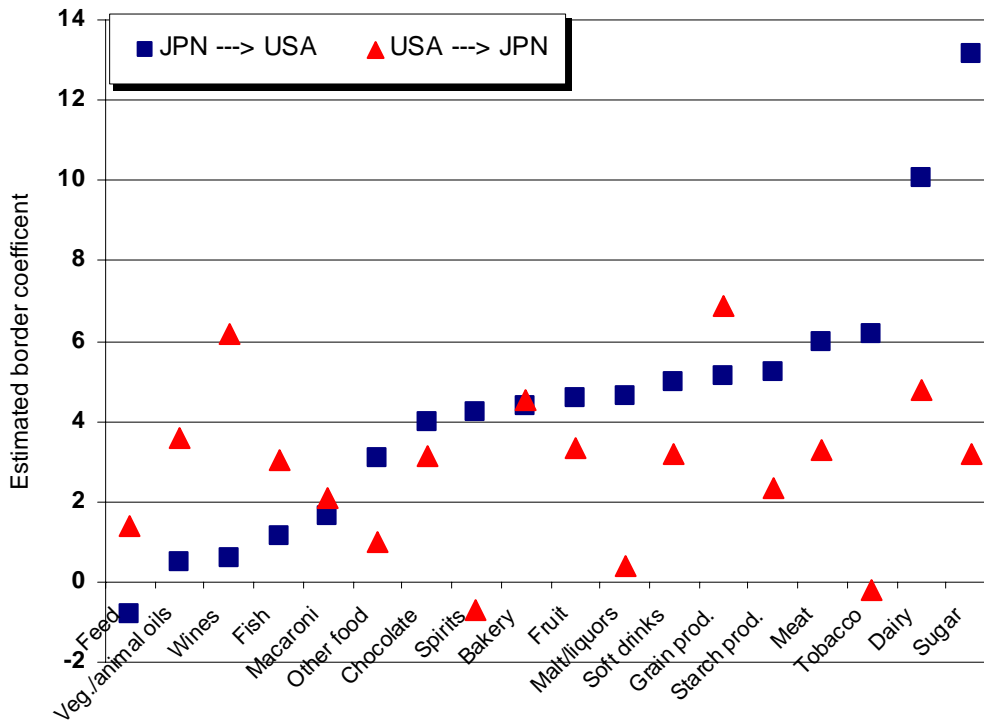


Figure 4: Market access asymmetry across industry between Japan and the US



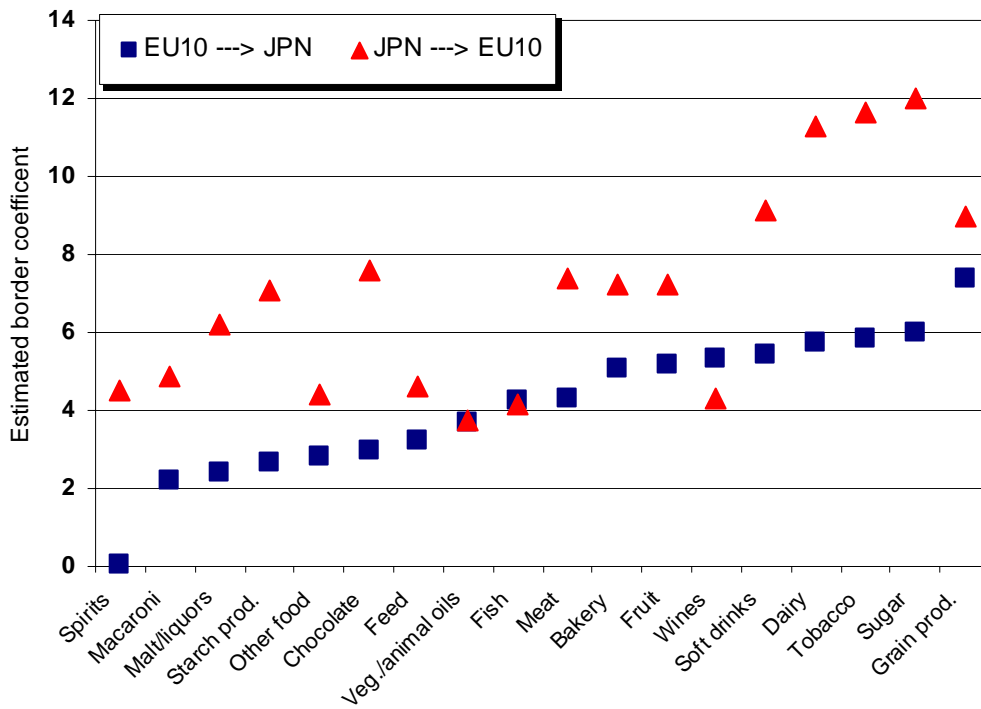


Figure 5: Market access asymmetry across industry between the EU10 and Japan

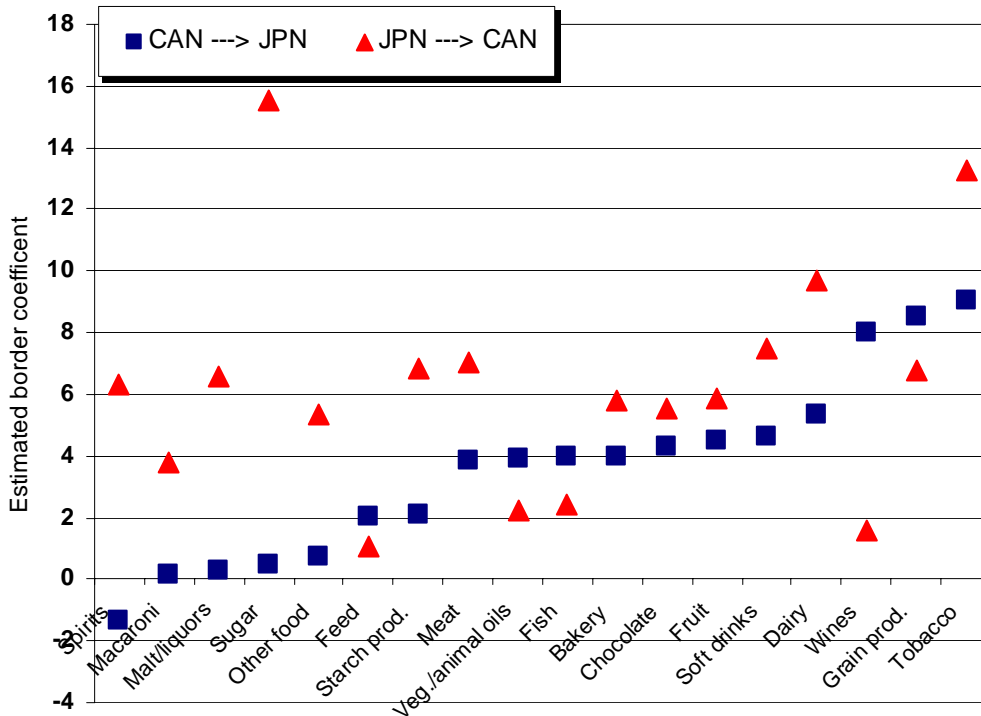


Figure 6: Market access asymmetry across industry between Canada and Japan