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# Preferences erosion and the developing countries exports to the EU: a dynamic panel gravity approach

***Valentina Raimondi*** *(University of Milano)* ***Margherita Scoppola*** *(University of Macerata)* ***Alessandro Olper*** *(University of Milano)*

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# Preferences erosion and the developing countries exports to the EU:

# A dynamic panel gravity approach[1]

Valentina Raimondi [2], Margherita Scoppola [3] and Alessandro Olper [2]

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**Abstract:** The 2003 reform of the Common agricultural policy implied a drastic change of the level and instruments of border protection in the rice industry. Because the EU grants trade preferences to a considerable number of developing countries exporting rice, the reform implied preferences erosion as well. This paper addresses the impact of preferences erosion on the rice exports of the countries to which the EU grants preferences, with the aim of contributing to the literature, from two main points of view: first, by proposing a new empirical approach to compute the preferential margin when tariff rate quotas are in force which is based on the assumption of the existence of fixed costs and economies of scale in the trading industry; second, by estimating the trade elasticities of preferences by means of a dynamic panel gravity equation to deal with the issue of endogeneity of the preferential margins and take into account persistency in bilateral trade flows due to sunk costs faced by the trading industry. Results show that the way preferential margins are calculated matters significantly when assessing the existence and extent of their erosion and the values of trade elasticities. Finally, the estimations highlight the fact that the impact of preferences is still very strong for some of the countries concerned.

**Keywords:** trade preferences, gravity model, GMM, tariff rate quotas, EU rice policy

**JEL code:** F13, Q17, F14

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[2] Dipartimento di economia e politica agraria, agro-alimentare e ambientale, Università degli studi di Milano

[3] Dipartimento di studi sullo sviluppo economico, Università di Macerata

**Preferences erosion and the developing countries exports to the EU:**

**A dynamic panel gravity approach**

## 1. Introduction

The erosion of preferences due to multilateral tariff reductions may result in significant export losses for developing countries. Multilateral liberalization reduces the competitive advantages of developing countries benefiting from trade preferences. Indeed, the reduction of Most Favored Nation (MFN) tariffs lowers the cost advantage of developing countries benefiting from preferences with respect to the other competitors. The erosion of preferences may exacerbate the problems developing countries often face in gaining access to markets in developed countries markets. An interesting case is the EU policy for rice imports. This policy has long been a consequence of domestic policy: both the level and the kind of trade protection were defined to guarantee the sustainability of domestic policy. After 2003, the reform of the Common agricultural policy implied a drastic change also of the level and instruments of border protection in the rice industry. The reform of domestic rice policy by the EU, which grants trade preferences to a considerable number of developing countries, by involving a reduction of the border protection, implied preferences erosion as well.

This paper focuses on the erosion of the preferences granted by the EU in the rice industry. Rice is among the most sensitive products for many developing countries exporting to the EU; for some of them, the EU represents a major export market and rice is among their most important export products. The objective of the paper is to assess the impact of the preferences erosion over the past decade on exports to the EU of developing countries benefiting from preferences and, more generally, to assess the current dependence of developing countries on EU preferences as regards their ability to access EU rice markets. For this purpose we use a gravity model. Compared to the previous literature estimating the trade impact of preferences by means of a gravity equation, this paper offers two main contributions. The first concerns the way in which the independent variable of interest, that is, the preferential margin, is calculated. As in other recent papers, the independent variable is a continuous – and not a dummy – variable (e.g. Cipollina, Salvatici 2010, Cardamone, 2009); further, our analysis is highly disaggregated and there is no bias due to tariff aggregation. Moreover, an innovative approach to calculate the preferential margin is proposed: since EU preferences to rice imports are granted by means of tariff rate quotas, to compute the preferential margin one needs to evaluate the actual tariff equivalent of the tariff rate quota. This paper proposes a new empirical strategy to calculate the tariff equivalent of a tariff rate quota, which is shown to be consistent with the assumption of fixed export costs and economies of scale in international trade. The paper compares the preferential margin obtained using this new approach with that obtained by the standard method showing that the latter may lead to a substantial underestimation of the preferential margin. The second contribution is the use of a dynamic panel gravity equation. As the literature has shown, the standard cross-section gravity model is unable to deal with endogeneity arising when estimating the trade preference effects, because of the difficulties in finding the appropriate instrumental variables (Baier and Bergstrand, 2007). Theoretically–based gravity models using panel data allow us to make adjustments for endogeneity due to omitted (selection) variable bias. Finally, the presence of exporter sunk costs raises the question of hysteresis and persistency in bilateral trade flows, an issue we deal with by estimating a dynamic version of the gravity equation, through a system-Generalized Method of Moment (GMM) estimator proposed by Blundell and Bond (1998).

Overall results show that the way preferential margins are calculated matters significantly when assessing the existence and extent of preferences erosion. Under the standard method there is no clear-cut evidence of preferences erosion, while the opposite is true when the tariff equivalent proposed in this paper is used. In the latter case, the results suggest that during the examined period there has been considerable erosion of preferences, even though the extent of the erosion varies across the different groups of countries. The method to calculate the margin also significantly affects the estimated values of trade elasticity, both in a static and a dynamic environment. More specifically, if there are fixed export costs and economies of scale, by using the standard tariff equivalent of tariff rate quotas one can significantly underestimate the (true) impact of preferences. Our estimations highlight the fact that the trade impact of preferences is currently still very high for some countries. Further, using the system-GMM estimator we estimate the short and the long run trade elasticities to preferences and the magnitude of the estimated long run coefficient confirms the inertial behaviour of exports.

The paper is organized as follows. The next section offers an overview of the EU trade policy in the rice industry. The third section explains the method used to calculate the tariff equivalent of tariff rate quotas and compares the preferential margins obtained with the standard approach with those obtained by using this new approach. The fourth section addresses the issues arising when estimating the trade impact of the preferences by means of the panel gravity equation, while the fifth illustrates the estimated models and the econometric strategy. The sixth discusses the results, while the final section draws various conclusions.

## 2. EU trade policy in the rice industry during the period 2000-2008: an overview

The international market of rice covers products that are rather diverse, from the point of view of both their characteristics and value added. Two main distinctive types of rice are traded - the Japonica and the Indica – and four different products: paddy, husked, milled and broken rice. Most EU imports are of husked (more than 60%) and milled rice (about 20%), while paddy rice imports are very small (less than 1%). Although the EU accounts for only 5,5% of world imports, it is a very important market for certain developing countries. For example, in 2007 the EU accounted for the 95%, 65%, 47% and 40% of the value of rice exports of Cambodia, Guyana, Bangladesh and Suriname, respectively.[4]

The EU trade policy in the rice industry is rather complicated; the instruments and the level of the border protection vary significantly across products and between imports regulated by multilateral agreements and those covered by the various preferential schemes. Before 2004, the tariffs applied to the EU imports on a MFN basis were defined by the 1994 GATT Agreement; while for paddy and broken rice specific fixed bound tariffs were applied, for husked and milled rice the applied tariff was established to be the smallest one between the bound tariff and the difference between a threshold import price and the international price. This threshold import price for the husked rice was equal to 180% (for the Indica rice) and 188% (for the Japonica rice) of the intervention price; for milled rice, it was set equal to the intervention price plus a percentage to be calculated. As a consequence of this import regime, tariffs applied to husked and milled rice fluctuated with the international price: when this was high, the tariff was the difference between the threshold import price and the international price and, hence, smaller than the bound tariff, but when the international price fell below a certain level then the bound tariff was applied.

With the 2003 reform of the Common Agricultural Policy the EU decided to reduce the value of the intervention price for rice drastically, by cutting it by 50%. The threshold import prices for husked and milled rice as well as tariffs consequently dropped. The EU and the main rice exporters then agreed to eliminate the threshold import price system and a new set of MFN bound tariffs for husked, milled and broken rice were negotiated, and entered in force in September 2004.[5] These new tariffs are significantly lower than the pre-reform values: in August 2004 the tariffs applied to imports were 197 Euro/t and 416 Euro/t for husked and milled rice, respectively, while in September 2004 these fell to 65 Euro/t and 175 Euro/t.

However, only 55% of EU imports of rice is currently subject to these MFN tariffs (COGEA, 2009).

A considerable amount of EU rice imports is currently covered by Tariff Rate Quotas (hereafter, TRQs), that is, a two-tiered tariff system with the volume imported within the quota charged at a lower tariff than out-of-quota imports. Several agricultural TRQs were introduced by the 1994 GATT Agreement on Agriculture to improve market access where agricultural protection was very high but, as regards EU rice imports, no TRQs were included in the Agreement. However, in application of the article XXIV of the GATT, after 1998 the EU granted a number of TRQs to the main rice exporters to compensate them for the 1995, 2004 and 2007 enlargements.[6] Country-specific TRQs were granted to the United States, Thailand,

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[4] These figures are drawn from COMTRADE database.
[5] While the value of the tariff applied to broken rice imports is fixed, for husked and milled rice three different values of tariffs may be applied depending upon the quantity imported. As for paddy rice, there was no need to set new tariffs as the threshold import price system was not in force in this case; hence, the applied tariff continues to be the 1994 GATT Agreement bound tariff.
[6] Hereafter, we will refer to these as the GATT TRQs.

Australia, India, Pakistan and Guyana for husked, milled and broken rice; further, there are also non-country specific TRQs. Imports under these GATT TRQs are estimated to account for about 30% of total EU rice imports in 2007 (COGEA, 2009)

Additional TRQs are granted by the EU under the preferential agreements (Table 1). In the rice industry, trade preferences are given exclusively by means of TRQs. Since the early Lomè Conventions a certain volume of rice from the ACP countries enters the EU at a lower tariff than the MFN one. More specifically, during the period examined in this paper, the EU granted a TRQ of 160.000 ton, 35,000 of which from the overseas countries and territories (OCT). In-quota tariffs were made up of two components: the first part is a percentage of the MFN tariffs, while the second is independent of the value of MFN tariff. Within the Generalized System of Preferences, Bangladesh benefits of a TRQ of 4,000 tons, with the in-quota tariff being made of two components as well. Under the Euro Mediterranean Agreement, the EU grants a TRQ of 32,000 tons to Egypt, with the in-quota tariff 25% lower than the MFN one. Finally, under the Everything But Arms initiative (EBA), a zero-duty TRQ has been in force since 2002, with the quota gradually increasing over the transitional period 2002-09. Almost 15% of total EU rice imports were covered by preferential TRQs in 2007 (COGEA, 2009).

## 3. Measuring preferential margins with Tariff Rate Quotas

### 3.1. The tariff equivalent of Tariff Rate Quotas

The presence of TRQs raises a number of issues when calculating preferential margins. One is finding the tariff equivalent of a TRQ. The literature on TRQs suggests that the tariff equivalent varies according to which of the three elements of a TRQ regime is binding (Boughner *et al.* 2000; Skully, 2001). Figure 1 illustrates the usual partial equilibrium framework under the assumptions of perfect competition and upward excess supply curve and three different excess demand curves. The excess supply curve is kinked: it is equal to $s+T^{in}$, where $T^{in}$ is the in-quota tariff, when imports are lower than the quota $\bar{Q}$; it is vertical when imports are equal to the quota; and is equal to $s+T^{out}$, with $T^{out}$ being the out-of-quota tariff, if imports are higher than the quota.

When import demand ($D_1$) is such that the equilibrium quantity is lower than the quota ($Q_1$), then the quota is not binding and the in-quota tariff is applied to all imports; in this case, the tariff that leaves imports and prices unchanged is clearly the one applied to the in-quota imports. In the second case the interaction between demand ($D_2$) and supply determines an equilibrium quantity ($Q_2$) higher than the quota; hence, there are out-of-quota imports. In this case, the out-of-quota tariff is applied to the out-of-quota imports, while the in-quota tariff to the in-quota imports. The equilibrium price is $P_2 = s(Q_2)+T^{out}$; the difference between the price $P_2$ and the marginal cost faced by traders importing within the quota, $s(\bar{Q})+T^{in}$ , is the unit rent caused by the quota. Clearly, the tariff that leaves imports and prices unchanged is the out-of-quota tariff. Finally, if demand ( $D_3$ ) crosses the supply curve on its vertical portion, the binding instrument is the quota itself. The value of the equilibrium price ( $P_3$ ) is in between $s(\bar{Q})+T^{in}$ and $s(\bar{Q})+T^{out}$ ; the difference between the equilibrium price and the marginal cost faced by importers ( $s(\bar{Q})+T^{in}$ ) is the unit rent. In this case the tariff equivalent is $P_3 - s(\bar{Q})$ .

The empirical literature relies on this theoretical framework to compute the tariff equivalent of TRQs. Many authors consider the in-quota tariff as tariff equivalent when imports are lower than the quota (case 1), the out-of-quota tariff when imports are higher than the quota (case 2) and an in-between value when imports are equal to the quota (case 3) (e.g. Cardamone, 2009; Garcia-Alvarez-Coque *et al*., 2010). Boumellassa *et al* (2009) determine the tariff equivalent of the TRQs on the basis of a range of fill rates in the database MAcMap-HS6v2. If the fill rate is lower than 90%, case 1 is adopted and, accordingly, the tariff equivalent is the in-quota tariff. When the fill rate is between 90% and 98% (case 3) the tariff equivalent is computed as the simple average of the in-quota and the out-of-quota tariff. Finally, if the fill rate is higher than 98% case 2 is adopted and the tariff equivalent is equal to the out-of-quota tariff.

The tariff equivalent of a TRQ may be different when economies of scale are considered. The usual framework used to analyze the economics of TRQs, illustrated in Figure 1, assumes that the excess supply curve of the exporting countries is upward sloping and, hence, that the marginal cost of importing agricultural goods is increasing. However, there are reasons to believe that this is not always the case. The costs faced by traders to import agricultural products include a variable component given, among other fasctors, by the cost of purchasing the agricultural good in the exporting countries. However, fixed costs are also often associated with international trading. These may arise because of the fixed costs traders sustain in acquiring knowledge of foreign markets; in addition, evidence exists that there are also economies of scale in shipping and in transportation in general (e.g. Hummels, Skiba, 2004).

To investigate the tariff equivalent of TRQs with fixed trading costs, we rely on the basic international trade model under economies of scale and monopolistic competition *à la* Dixit-Stiglitz-Krugman (see Feenstra, 2003).[7] In this setting, a number of (symmetric) firms are assumed to produce differentiated products; each firm is a monopolist for the variety it produces and, thus, it maximizes profits by equalizing marginal revenues with marginal costs; marginal costs are assumed to be constant. Because of fixed costs, the average cost declines with imports and is always higher than the marginal cost; as each firm's profits are positive, if there are no restrictions to entry, new firms enter the market. This reduces the market share of each firm and increases the average cost; in equilibrium, profits are zero and the price equals the

[7] The importance of fixed costs in international trade has been recently emphasized by the firm-level heterogeneity literature (e.g. Melitz, 2003, Jorgenson, Schroeder, 2008). For simplicity, the framework here assumes symmetry both on the demand and supply side.

average cost. Because of the assumption of symmetry, prices and quantities are identical across all varieties; the price and the imported quantity of the variety $i$ are thus also the price and quantities of all imported varieties.

Figure 2 illustrates the analysis of the TRQ under these assumptions. The average cost, $AC$ , of the importing firm under free trade is:

$$AC = \frac{FC}{Q} + c \tag{1}$$

where $FC$ are the fixed cost, $Q$ is the imported quantity and $c$ is the constant variable cost.

If, as above, $\bar{Q}, T^{in}$ and $T^{out}$ are the quota, the in-quota and the out-of-quota tariffs, respectively, then under the TRQ the average cost is:

$$AC_{T^{in},T^{out}} = \begin{cases} \dfrac{FC}{Q} + c + \dfrac{T^{in}\bar{Q} + T^{out}(Q-\bar{Q})}{Q} & \text{if } Q > \bar{Q} \\ \dfrac{FC}{Q} + c + T^{in} & \text{if } Q \leq \bar{Q} \end{cases} \tag{2}$$

In equilibrium, the price is equal to $AC_{T^{in},T^{out}}$ . The Figure reports two demand curves faced by the monopolistic firm under equilibrium, which reflect different market sizes. As market size increases the firm can exploit economies of scale, thereby incurring lower average costs; positive profits attract new firms and this increases the degree of competition on the market and the elasticity of the demand faced by each firm. Thus, the larger the size of the market, the higher the elasticity of the demand faced by each firm. $D_1$ is the demand curve when the market size is small, relative to the quota; the equilibrium quantity is $Q_1 < \bar{Q}$ and the price under the TRQ is $P_1$. Clearly, the tariff that leaves unchanged the price and the imported quantity is the in-quota tariff $T^{in}$ , such as under perfect competition and increasing costs. However, if the market size is large enough with respect to the quota ( $D_2$ ), then the equilibrium quantity, $Q_2$ , is higher than the quota and the equilibrium price is $P_2$ . In this case, the tariff which would leave price and imports unchanged is the weighted average of the two tariffs. Finally, when the demand curve crosses $AC_{T^{in},T^{out}}$ for $Q = \bar{Q}$ , the tariff equivalent is again the in-quota tariff $T^{in}$ .

Hence, within this framework if imports are no greater than the quota, the tariff equivalent is the in-quota tariff; alternatively it is the weighted average of the two tariffs. Thus, the tariff equivalent computed on the base of the economies of scale-monopolistic competition framework is always no greater than the one consistent with the perfect competition model.

### 3.2. Preferential margins of rice exporters to the EU: a comparison of different approaches

To compare the preferential margins (PM) computed under different hypotheses, a data base has been built which includes the applied in-quota and out-of-quota tariffs and the quantities imported within the quota and out-of-the quota. The database covers 36 rice products (HS-8 digit level) and 123 producing and/or exporting countries for 9 years (2000-08). By using tariffs at a highly disaggregated level there are no distortions due to tariffs aggregation, as EU tariffs in the rice industry are defined at the HS-8 digit level. [8] Data on actual imports within the quota are not easily available and, thus, many studies calculate in-quota imports by comparing the granted quota with total imports (e.g. Cardamone, 2009; Garcia-Alvarez-Coque *et al*., 2010). If total imports are equal to or exceed the quota, in-quota imports are set as equal to the quota; alternatively, in-quota imports are equal to total imports. In this way, one is implicitly assuming that the quota is filled. However, evidence about the fill rate of TRQs suggests that usually the opposite is true, that is, the fill rate is seldom equal to 100% (WTO, 2006). In this paper, in-quota imports are directly drawn from the EU Commission, which collects the amount of product that has been actually imported within the TRQs, at the HS-8 digit level. By using the real amount of product imported at the in-quota tariff, no *a priori* assumption about the fill rate of the quota is made.

As out-of-quota imports are not collected by the EU Commission, they are computed here as the difference between total yearly imports from the Comext database, and the in-quota imports data collected by the EC Commission. In this way, however, out-of-quota imports can be slightly overestimated or underestimated. This is because in-quota imports provided by the EC Commission are registered in the year in which licenses are issued, while Comext data refer to the year in which the product actually enters the EU. As licenses are valid for a few months, it is possible that in-quota imports registered for a certain year actually enter in the EU in the following year.[9] This potential error in calculating the out-of-quota imports may in principle have significant implications for the tariff equivalent of TRQs under the assumption of perfect competition: in this case, very small errors in the out-of-quota imports may lead to serious errors in assessing the value of the tariff equivalent.

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[8] Tariffs have been converted in *ad valorem* tariffs by using import unit values, given by the ratio between the value and the quantity of the EU imports for each product and each year.

[9] To make this point clearer, consider the following example: assume that in 2004 and 2005 the licenses allocated by the EC are 100 per year, which are fully used in both years, and that traders imports 2 of the 2004 licenses in the first weeks of 2005, but do not import out-of-the quota. Comext data indicate that imports are 98 in 2004 and 102 in 2005; thus, we conclude that there are 2 out-of-quota imports in 2005. Under the assumption of perfect competition, the tariff equivalent would be the in-quota tariff in 2004 and the out-of-quota tariff in 2005, which are both obviously incorrect, because in both years imports are equal to the quota and, thus, the tariff equivalent is always in-between the two tariffs.

If $T_{kj}^{PREF}$ is the preferential *ad valorem* tariff and $T_{k}^{MFN}$ is the MFN *ad valorem* tariff, with $k$ and $j$ being the product and the exporting country, respectively, the general formula used to calculate the preferential margin in a certain year is the following:

$$PM_{kj} = \frac{T_{k}^{MFN} - T_{kj}^{PREF}}{1 + T_{kj}^{PREF}} \tag{3}$$

Two different PM have been computed to take into account the two alternative measures of the tariff equivalent. Under the perfect competition hypothesis, if $Q_{kj}$ are total imports and $\bar{Q}_{kj}$ is the quota, in a given year the PM is the following:

$$PM_{kj}^{P} = \left\{ \begin{array}{ll} \dfrac{T_{k}^{MFN} - T^{in}{}_{kj}}{1 + T^{in}{}_{kj}} & \text{if } Q_{kj} < \bar{Q}_{kj} \\ \dfrac{T_{k}^{MFN} - T^{out}{}_{kj}}{1 + T^{out}{}_{kj}} & \text{if } Q_{kj} > \bar{Q}_{kj} \\ \dfrac{T_{k}^{MFN} - \dfrac{(T^{out}{}_{kj} + T^{in}{}_{kj})}{2}}{1 + \dfrac{(T^{out}{}_{kj} + T^{in}{}_{kj})}{2}} & \text{if } Q_{kj} = \bar{Q}_{kj} \end{array} \right\} \tag{4}$$

It is worth noting that the tariff $T_{kj}$ applied to imports exceeding the preferential TRQs may be lower than $T_{k}^{MFN}$, because the EU may also grant the (preferred) exporting country TRQs within the GATT. For example, Egypt exports broken rice to the EU within preferential TRQs, but there are also additional imports which are charged at the in-quota tariff of the GATT TRQs.

The preferential margin under the assumption of economies of scale is:

$$PM_{kj}^{E} = \left\{ \begin{array}{ll} \dfrac{T_{k}^{MFN} - T_{kj}^{in}}{1 + T_{kj}^{in}} & \text{if } Q_{kj} \le \bar{Q}_{kj} \\ \dfrac{T_{k}^{MFN} - \dfrac{(T_{kj}^{out}(Q_{kj} - \bar{Q}_{kj}) + T_{kj}^{in}\bar{Q}_{kj})}{Q_{kj}}}{1 + \dfrac{(T_{kj}^{out}(Q_{kj} - \bar{Q}_{kj}) + T_{kj}^{in}\bar{Q}_{kj})}{Q_{kj}}} & \text{if } Q_{kj} > \bar{Q}_{kj} \end{array} \right\} \tag{5}$$

As for the tariff $T_k^{MFN}$, the maximum tariff applied to the $k$ product across all non-preferred exporters has been considered as the relevant MFN tariff.

Table 2 reports the different values of the PM computed for EU imports of husked rice from Guyana, which well illustrate how the assumptions made on the value of the tariff equivalent of the TRQ~~,~~ may affect the value of the margins. Margins are also reported in absolute terms, that is, by considering only the numerator in (4) and (5). The first column shows that in five out of nine years Guyana exported out-of-the preferential quota. Data confirm, as expected, that $PM^E \geq PM^P$. When out-of-quota imports are zero, the tariff equivalents computed under the two different hypothesis are identical and, thus, $PM^E = PM^P$. However, when there are out-of-quota imports, the tariff equivalent consistent with the assumption of perfect competition is higher and the margin is lower. As the Table shows, even a small amount of out-of-quota imports, as in 2001, may sharply reduce $PM^P$. Overall, $PM^E$ indicates that preferential margins before 2004 ranged between 18% and 25% while after 2004 they drop to less than 10%, thus confirming the existence of remarkable preference erosion following the policy reform of 2004.[10] The evidence is less clear-cut for the values of $PM^P$, as in four out of nine years this is equal to zero because of positive out-of-quota imports.

PM have been also aggregated by product and by country through the weighted averages of the $PM_{kj}$, with the weights being the imported volume in the whole period of a certain product/year.

Figures 3 and 4 show the evolution of the PM for three HS-6 digit products under the hypothesis of economies of scale and of perfect competition, respectively. After 2004, $PM^E$ sharply decreased especially for milled and husked rice. For husked rice, $PM^E$ fell from an average value of 22% in the period 2000-04 to an average of 9% in the 2005-08 period, while for milled rice the reduction was from 33% to 13%. This suggests considerable erosion of the preferential margins after the policy reform of 2004. $PM^E$ of broken rice declined as well after 2004, albeit to a lower extent.

Evidence of preferences erosion is less clear under the assumption of prefect competition (Figure 4). As mentioned above, $PM^P$ may vary enormously from one year to the next, because a small amount of out-of-quota imports implies a collapse of the $PM^P$. This occurs, for example, for husked rice in 2003 and 2004, right before the reform, or for milled rice in 2000. As for husked rice, $PM^P$ decreased from an average value of 10% in the period 2000-04, to an average of 4% in the 2005-08 period, while for milled rice the reduction was from 13% to 6% (the average $PM^E$ from 33% to 13%). Overall, by using $PM^P$ the extent of the preference erosion after 2004 is significantly lower: by observing $PM^P$ one concludes that PM has

[10] In-quota tariffs granted to ACP countries since 2003 have slightly declined (Table 1); the drop in the margin is therefore entirely explained by the fall in the MFN tariffs.

reduced by 6 and 7 percentage points for husked and milled rice, respectively; but these reductions are considerably higher if the $PM^E$ is taken into account (13 and 20 percentage points for husked and milled rice, respectively).

Figures 5 and 6 show the average PM by group of preferred countries. The values of $PM^E$ indicate that the margins after 2004 clearly declined for all groups, with the EBA countries showing the sharpest decline. This may be explained by the different way in which the EU grants preferences to the ACP compared to the EBA countries. The value of the preferred tariffs granted to the ACP countries is partly linked to the value of the MFN tariff (Table 1); as a consequence, the considerable reduction of the MFN tariffs after 2004 was not fully transmitted to the PM, because the preferred tariffs also diminished, albeit to a lesser extent. On the contrary, EBA countries during that period benefited from a zero in-quota tariff; hence, the reduction of the MFN tariffs was wholly translated into a reduction of the PM. Egypt benefitted from lower preferences than EBA and ACP countries preferences. [11] But in this case the fall in the PM is not due to the fall in the MFN tariffs, because the preferential tariff was defined as a percentage of the value of the MFN tariff; thus, the former declined proportionally with the latter, implying only negligible changes in the value of the PM. In fact, Egypt's PM drastically declined in the final years of the period because Egypt started to export considerable amounts of broken rice out-of-the-quota at MFN tariffs.

Once again the values of the $PM^P$ (Figure 6) for the three group of countries provide no clear indication of erosion after 2004; for example, there is no clear-cut evidence of preference erosion in EBA countries, mainly because in certain years they imported small quantities out-of-the quota, even if their TRQs were not wholly filled. [12] This occurred in 2002, 2004, 2005 and 2006. Hence, the $PM^P$ becomes zero in three years and almost zero in 2005. It is well known that least developing countries often face difficulties in exploiting preferences, because requesting preferences is a costly procedure especially when a quota is in place. The evolution of the $PM^P$ indicates that there was no preference erosion because the least developing countries were able to import (even if in small amounts) out-of-the quota before and after 2004. When observing the $PM^P$ one can find no clear-cut evidence of preference erosion for the ACP countries either; this sharply declined in 2003 because of out-of quota imports which occurred despite the TRQ being unfilled, while in 2002 there were no out-of-quota imports and the margin was rather high. Overall, because the ACP and the EBA countries never filled their TRQs, the fluctuation in the $PM^P$ reflects the ability of countries to use preferences and this varies from year to year, as seen in the different values of the TRQ fill rate over the period. On the basis of the $PM^P$ one should conclude that there was no erosion of preferences after 2004. The $PM^P$ indicates that preferences to Egypt dropped to zero after 2004 but, as mentioned

[11] In-quota tariff in this case was equal to the 75% of the MFN tariff (Table 1), which is much higher than the tariffs granted to the ACP and to EBA countries.
[12] As for EBA countries, the fill rate ranged between 56% to 79% over the period.

above, this is not due to the 2004 reduction of the MFN tariffs, but rather to Egypt's improved ability to export out-of-the quota at the MFN tariffs.

### 4. Estimating the trade effect of preferential margins with gravity equation

The preferential margins determined on the basis of the two different approaches have been used in a gravity equation to estimate the trade impact of preferences. The literature studying the average treatment effect of trade preferences using the gravity equation is largely based on the assumption that $PM$ is an exogenous variable (see, e.g., Cipollina and Salvatici, 2010; Nilsson and Matsson, 2009; Cardamone, 2009). This approach consistently identifies the average treatment effect of $PM$ if the economic agents' decision to select a programme is unrelated to unobservable factors influencing the outcome. However, as discussed in Bair and Bergstrand (2004; 2007), in the context of free trade agreements (FTA), many trade-policy analysts have noted that trade inhibiting policies, such as non-tariff barriers, may be one of the main reasons why governments select a specific FTA.

In this specific context we face a similar problem. Indeed, the EU decision to adopt a preferential regime could also be, among other things, a function of several unobservable factors: for example, the existence of specific domestic regulations, such as the stringency of the EU food safety and quality standards, as well as political motives unrelated to trade. In this context, countries select a preferential regime for reasons that are difficult to observe and are often correlated with the level of trade. This raises the classical problem of endogeneity in RHS variables.

Endogeneity usually arises under three forms: omitted variables, measurement error, and simultaneity bias (Wooldridge, 2002). While the use of a continuous instead of a dummy variable to measure the preferences can mitigate *measurement error* bias, Baier and Bergstrand (2007) suggest that *omitted variable* (selection) bias and, to a lesser extent, simultaneity remain the major sources of endogeneity in the estimation of the effects of trade preferences by means of the gravity equation. In this situation, the standard cross-country gravity equation is unable to account for endogeneity, as any potential instrument for trade preferences is also a determinant of bilateral trade (Magee, 2003; Baier and Bergstrand, 2004). The most plausible estimate of the average effect of an FTA, that allows us to account for endogeneity due to omitted variable bias, is obtained from (theoretically-based) gravity models using panel data (Baier and Bergstrand 2007, Magee 2008, Martinez-Zarzoso *et al*. 2009).

Specifically, the panel gravity equation should include time-varying country dummies to account for time-varying multilateral-resistance terms as well as to eliminate the bias from the gold-medal error identified by Baldwin and Taglioni (2006). In this way, variables that are difficult to measure, such as "infrastructure, factor endowments, multilateral trade liberalization, and unobserved time-specific shocks, will be captured by the *importer-year* and *exporter-year* fixed effects." (Magee, 2008 p. 353). Last but not least, the presence of unobserved time-invariant bilateral factors simultaneously influencing the presence of an FTA and the volume of trade have to be controlled for by *country-pair* fixed effects (Baier and Bergstrand, 2007).

To estimate the average effect of the $PM$ on rice exports to the European Union we follow this strategy although, unlike previous contributions estimating the trade effect of preferential agreements in a panel data setting, we use a continuous, instead of a dummy, preference variable in order to evaluate how this average effect changes with the use of different methods to calculate the $PM$ .

As mentioned in section 3, trade flow data come from the External trade statistics (Comext), produced by Eurostat which provides the value and the quantity of goods traded by EU member states with third countries. Due to the common nature of the EU trade policy, the EU is treated as a single entity; hence, we consider aggregated EU imports from all sources, as well as take into account the enlargement processes in 2004 and 2007. As for the dependent variable, we take account of overall trade, and not just that benefiting from preferences, as was often the case in previous papers (see, e.g.,Nilsson and Matsson, 2009). Indeed, there are several reasons that call into question the use of preferential trade only, that are related to both spill-over effects and the reallocation of market shares towards more productive firms. First, when a firm decides to export to the EU after the introduction of a preferential tariff – for rice, this was the case with the zero-duty quota introduced in 2002 under the EBA initiative – it has to face sunk costs linked to the marketing of the product, such as the new (trade) infrastructures and transaction costs to meet the EU standards and, eventually, the setup of a foreign distribution chain (Arkolakis, 2008). These may generate spill-over effects on total trade, as they are likely to improve the overall ability to export to the EU. Second, as suggested by the recent trade theory, exposure to international trade induces the more productive firms to export while simultaneously forcing the least productive firms to exit. Both the exit of the less productive firms and the additional exports sales gained by the more productive firms reallocate market share towards the latter (Melitz, 2003). As a consequence of this, the ability of the average firm to export increases irrespective of the existence of preferences. Finally, this productivity boost of exporting firms is also attributable to the effect of the learning process (Greenaway and Kneller, 2007) that will clearly affect trade overall and not just preferential trade.

This leads us to the issue of *persistency* and *hysteresis* in bilateral trade which need to be accounted for in the empirical analysis. Indeed, even when the original reason for a high level of bilateral trade has disappeared, the stock of capital that firms have invested in the form of marketing and distribution networks, brand-name loyalty among customers, and so forth, live on for many years thereafter. The word *hysteresis* is sometimes applied to this phenomenon, suggesting that the effect is considered to be permanent (Baldwin, 2006). A set of theoretical models by Dixit (1989), Krugman (1989), and others suggest that hysteresis in exports may be due to sunk costs in entering the export market at the firm level. Thus in order to tackle *hysteresis* in trade, we have estimated the gravity equation dynamically. This approach allows us to distinguish between the short-run and the long-run impact, with the latter capturing the observed evidence that countries trading with each other tend to have an inertial behaviour, possibly due to sunk costs.

## 5. Empirical specification of the gravity equation

### *5.1 Static gravity equation*

The standard gravity equation commonly estimated using cross-section data is:

$$m_{ijk} = \delta_0 \left(GDP_i\right)^{\delta 1} \left(GDP_j\right)^{\delta 2} \left(d_{ij}\right)^{\delta 3} \left(t_{ijk}\right)^{\delta 4} e^{\delta 5\left(Lang_{ij}\right)} e^{\delta 6\left(Cont_{ij}\right)} \varepsilon_{ijk} \qquad (6)$$

where $m_{ijk}$ is the trade flow to country *i* from country *j* of good *k*; $GDP_t$ ($GDP_j$) is the nominal gross domestic product in the destination (origin) country; $d_{ij}$ reflects the impact of transport costs and is proxied by distance between countries; *Lang* and *Cont* are binary variable assuming the value 1 if *i* and *j* share a common language or a common border, and 0 otherwise. Finally, $t_{ijk}$ are the trade policies, proxied by the *ad valorem* equivalent tariff factor imposed by country *j* on commodity *k* imports from country *i*:

$$t_{ijk} = \left(1 + T_{ijk}\right)$$

with $T_{ijk}$ being the *ad valorem* equivalent tariff. Rewriting equation (6) in logarithmic form and introducing the time dimension, as well as the fixed effects in accordance with the theory, the basic empirical model can be expressed as:

$$\ln m_{ijkt} = \beta_0 + \beta_1 \ln\left(1 + T_{ijkt}\right) + \alpha_{jt} + \alpha_{it} + \alpha_{ij} + \alpha_{hs6t} + \alpha_t \qquad (7)$$

where $\alpha_{ij}$ are bilateral fixed effects to control for heterogeneities accounting for the impact on trade of any observed and unobserved characteristic of a country pair that is constant over time, such as the distance between countries (proxy of transportation costs), a common language or common border, colonial relationship as well as other historical, cultural and political ties between trading partners (Magee, 2008); $a_{it}$ and $a_{jt}$ are the importer-year and exporter-year fixed effects that account for country variation in real GDP, population as well as other difficult to measure variables such as infrastructure, factor endowments or time specific shocks. These country-and-time effects account explicitly for the time-varying multilateral price terms (Anderson and van Wincoop, 2004; Baier and Bergstrand, 2007). Finally $\alpha_t$ and $\alpha_{hs6t}$ are year and product-time dummies to account for any shocks that affect global trade flows in a particular year or in a particular time-product group, respectively.

In this specific case, because we consider the EU as the unique importer, the importer-year $\alpha_{it}$ and bilateral fixed effects $\alpha_{ij}$ are dropped because they are perfectly collinear with the time dummies and the exporter-year dummies. Moreover, our definition of $PM$ (see equation (3)) can be written as: $(1+T_{kjt}^{PREF}) = (1+T_{kt}^{MFN})/(1+PM_{kjt})$ .

Plugging this relation into equation (7) we obtain:

$$\ln m_{jkt} = \beta_0 + \beta_1\left[\ln\left(1+T_{kt}^{MFN}\right) - \ln\left(1+PM_{jkt}\right)\right] + \alpha_{jt} + \alpha_{hs6t} + \alpha_t \qquad (8)$$

Finally, since $T_{kt}^{MFN}$ does not vary across exporters, it is fully captured by time-product fixed effects, thus the final static panel gravity specification becomes

$$\ln m_{jkt} = \beta_0 + \beta_2 \ln\left(1+PM_{jkt}\right) + \alpha_{jt} + \alpha_{hs6t} + \alpha_t \,. \qquad (9)$$

To estimate equation (9) consistently we follow the standard practice in gravity literature (see Martin and Pham, 2008; Helpman *et al.*, 2008) of implementing the Heckman two stage selection correction procedure (Heckman, 1979). In a panel data setting, this means to estimate a panel random-effects Probit equation with exporter and importer fixed effects and time effect, as first step selection equation. From this estimation, the inverse Mill ratio is retrieved and included as regressor in the so-called output equation, namely a least square regression with dummy variables (LSDV) that include time and exporter-year dummies (see Martinez-Zarzoso *et al.* 2009).

Finally, to check for robustness we also applied an alternative approach using the Poisson Pseudo Maximum-Likelihood (PPML) estimator proposed in the influential paper of Silva and Tenreyro (2006) to solve heteroscedasticity problems in the gravity model.[13]

*5.2 Dynamic gravity equation*

To account for persistency and hysteresis in trade flows equation (9) could be specified dynamically by adding the lagged dependent variable on the right-hand side.

$$\ln m_{jkt} = \gamma_0 + \gamma_1 \ln m_{jk(t-1)} + \gamma_2 \ln\left(1+PM_{jkt}\right) + \alpha_{jt} + \alpha_{hs6t} + \alpha_t + u_{jkt} \;. \qquad (10)$$

where $\gamma_1$ is the adjustment coefficient in the dynamic model.

The introduction of dynamics raises econometric problems when the time span of the panel is short, as in our application. Indeed, the correlation between the lagged dependent variable and the transformed

[13] Martin and Pham (2008) have shown that the Heckman method performs better if true identifying restrictions are available. Conversely, the PPML solves the heteroscedasticity problem, but yields biased estimates when zero trade observations are frequent.

error term renders the least squared within estimator biased and inconsistent in panels with large cross-sections and short time series. To avoid this inconsistency, Arellano and Bond (1991) proposed a Generalised Method of Moments (GMM) estimator as an alternative to LSDV. They suggested transforming the model into a two step procedure based on first difference to eliminate the fixed effects, as a first step. In the second step, the lagged dependent variable is instrumented using the two period lagged differences (or two period lagged level) of the dependent variable.[14]

In the case of the gravity model, first-differencing the equation removes the fixed effect but also the time invariant regressors of the specification and, when the regressors are of interest, the resulting loss of information may be a serious drawback (De Benedictis and Vicarelli 2005). Moreover, with highly persistent data and short panel (along the time dimension), as in the case of all bilateral exports flows, and of our dataset specifically, the GMM estimator may suffer marked small sample bias due to weak instruments (Blundell and Bond 1998).

As a solution, Arellano and Bover (1995) and Blundell and Bond (1998) built a system of two equations, known as System-GMM, which supplements the equations in first differences with equation in level. In particular, the System-GMM estimator utilises instruments in level for the first-differenced equation and first-differenced instruments for the equation in level. Following the Blundell and Bond system equations, the gravity specification is:

$$\begin{aligned} d\ln m_{jkt} &= \gamma_0 + \gamma_1 d\ln m_{jk(t-1)} + \gamma_2 d\ln\left(1+PM_{jkt}\right) + \\ &+ \gamma_3 d\ln v_{jt} + \alpha_{jt} + \alpha_{hs6t} + \alpha_t + u_{jkt} \end{aligned} \tag{11}$$

and

$$\begin{aligned} \ln m_{jkt} &= \gamma_0 + \gamma_1 \ln m_{jk(t-1)} + \gamma_2 \ln\left(1+PM_{jkt}\right) + \gamma_4 \ln v_{jt} + \\ &+ \gamma_5 \ln dist_j + \alpha_{jt} + \alpha_{hs6t} + \alpha_t + u_{jkt} \end{aligned}, \tag{12}$$

where $d$ denotes first differences, $v_{jt}$ the exporter rice production volume and is treated as predetermined; $dist_j$ the distance between the exporting country and the European Union, considered as a strictly exogenous covariate;[15] and finally, the lagged dependent variable $m_{jk(t-1)}$ and the preferential factor ($1+PM_{jkt}$) that are treated as endogenous. Thus, the GMM estimator also represents a natural strategy to account for the endogeneity of the preference factor, as well as measurement error and weak instruments, while controlling for time-invariant country specific effects such as distance.

[14] All runs using the Hansen (1982) two-step GMM estimator.
[15] Rice production volumes come from FAOSTAT database, while distances between countries come from CEPII database.

Following Martinez-Zarzoso *et al.* (2009), we consider that by including lagged bilateral exports in the right hand side of the equation we are able to control for the time-varying components of the multilateral resistance term. Consequently, neither time-varying exporter dummies nor other explicit fixed effect dummies are included in the GMM regressions, (see Roodman, 2009b).

## 6. Econometric results

### *6.1 Static model results*

We start by estimating a cross-section gravity equation for single years of the time period covered. Table 3 provides the preferential margin impact for the years 2001, 2005, and 2008. The estimated coefficient of interest is $\beta_2$, that represents trade elasticity to the factor margin ($1+PM_{jkt}$). The two sets of estimates, for both the 'Standard' margin $PM^P$ and the 'weighted' margin $PM^E$, that accounts for economies of scale and imperfect competition, present quite unstable coefficients from year to year and in some years are even negative as regards $PM^P$. With a value of about 14, the only statistically significant elasticity estimate is that related to 2008, and refers to $PM^E$. Thus, it appears quite difficult to reach any conclusion about the effect of $PM$ on trade flows from these cross-section results.[16]

While several reasons can be put forward to explain this instability, the preliminary evidence confirms the recent literature that criticises the use of cross-section regressions to infer the average effect of preferential margins (Baier and Bergstrand, 2007; Martinez-Zarzoso et al. 2009). Indeed, as discussed above, the simple inclusion of country fixed effects does not correct the endogeneity bias caused by the country selection of preferential regimes. In a cross-section gravity equation, we should use IV technique to correct this endogeneity bias. However, finding good instruments correlated with $PM$ and uncorrelated with bilateral trade is a well known problem in the gravity literature.

Econometric evidence based on panel data is reported in Table 4. Columns 1-2 present regression results when the gravity model is estimated over the time period 2000-2008, using LSDV with country-time fixed effects. Column 1 includes the 'Standard' margin $PM^P$, while column 2 considers the 'weighted' margin $PM^E$. Under perfect competition, the trade elasticity of the preferential margin factor in the rice sector, namely its estimated coefficient, has a magnitude near to 5. Interestingly, the effect of the estimated preferences clearly increases in magnitude when the $PM^E$ is considered. Specifically, the coefficient more than doubles, passing from 4.9 to 11.4.

Columns 3 and 4 estimate the preference effects taking into account problems of selection bias and thus adding to the second step Heckman equation the inverse Mills ratio, retrieved from the first step (probit)

[16] The instability of coefficients of Table 3, obtained from the non-zero trade flows only, are generally unaffected by the use of the Heckman procedure to control for sample selection (results not reported).

selection equation.[17] The marked presence of zero trade in our dataset (about 80%) makes the inverse Mill ratio significant, giving evidence of selection bias. Both $PM^P$ and $PM^E$ coefficients notably increase in magnitude, and this is particularly true for $PM^P$. Indeed, the magnitude of the estimated effect of $PM^E$ is now only slightly higher than $PM^P$.

To check for robustness, columns 5-6 of Table 4 report estimates of the gravity equation using the PPML estimator.[18] The trade elasticities are consistently higher than the LSDV ones, and (as expected) quite close to those obtained with the Heckman procedure, confirming the importance of sample selection in the dataset. More importantly, also the PPML results display a trade elasticity of $PM^E$ significantly higher than that of $PM^P$ (18.5 vs. 10), reinforcing our key finding about the bias introduced in the TRQ tariff equivalent estimation when perfect competition is assumed.

Thus, whatever the estimation method, the message is similar: assuming scale economies and imperfect competition (vis-à-vis perfect competition) to measure the TRQ tariff equivalent significantly increases the sensitivity of trade flows to preference margins. Thus, the preliminary evidence suggests that the way the tariff equivalent of TRQs is computed is crucial for the estimation of the trade preference elasticity, reinforcing our stylized facts reported in section 3.

However, we need to deal with a further econometric problem. While the panel gravity specification reported in Table 4 accounts for several problems highlighted by the recent gravity literature and, especially, the endogeneity of *PM*, it fails to control for potential persistency in trade flows due to fixed export costs, an issue that is at the root of our modelling approach. Thus, in the next section we shall focus our attention on this potential source of bias by estimating a dynamic panel gravity model.

### *6.2 Dynamic model results*

For comparative purposes we start by estimating the dynamic model with the classical LSDV specification. Columns 1 and 2 of Table 5 present the results when the model is estimated with the within fixed effects estimator that includes the lagged exports, while Columns 3 and 4 report the results obtained with the system-GMM estimator. Both methods distinguish the 'Standard' margin $PM^P$ and the 'weighted' margin $PM^E$ .

Using the LSDV estimator, the coefficients of the lagged exports are positive and significant, supporting the idea that the gravity model should be estimated in a dynamic panel setting. The estimated (short-run) trade elasticity to preferences is positive for both $PM^P$ and $PM^E$, but only the latter is significant

[17] The probit selection equation (not reported) presents estimated coefficients that are statistically significant and with the expected signs. As expected, $PM$ increases the probability of registering positive trade flows.
[18] In the PPML procedure, we used product dummies instead of time-product dummies due to convergence problems from the high number of dummies. Results obtained using a smaller sample show tiny variations in the estimated coefficients.

at 5% level, with a value of around 4.3. However, as in the LSDV estimator, the lagged variable is correlated with the fixed effects in the error term, this estimator does not eliminate the bias (Roodman, 2009a).[19]

The coefficients of the lagged exports estimated through the system-GMM proposed by Blundell and Bond (1998) are, as expected, positive, statistically significant and particularly high (0.61), confirming that countries trading heavily with each other are expected to continue to do so. The magnitude of the persistence effect is quite similar to previous findings (see, e.g., Vicarelli *et al.* 2008, Martinez-Zarzoso et al. 2009), especially when the data used are highly disaggregated. The bottom of the Table reports the AR(2), the Hansen tests and the difference-in-Hansen tests to check the consistency of the GMM estimator, the lack of autocorrelation of the residuals and the validity of the instrumental variables. The Arellano-Bond test for autocorrelation AR(2) indicates that second order correlation is not present.[20] Moreover, the standard Hansen test confirms that in all cases our set of instruments is valid (difference-in-Hansen checks the validity of a subset of instruments). Moreover, according to Roodman (2009a), the instrument count does not exceed the number of groups and, to control for instrument proliferation that could cause a weak Hansen test, we used only 3 lags instead of all available lags for instruments.

The estimated coefficients from the GMM-system are interpretable as in a standard linear model. According to the standard gravity model, the coefficients of distance and production present the expected sign, but the former is not significant and the latter is relatively small. The last result is probably due to the fact that production is here proxied by quantities, and not by values as required by gravity theory.

In line with results reported in the previous section, we find that the trade elasticity of the preferential margin factor $(1+PM)$, under the assumption of perfect competition, is never significant. In contrast, the preferential margin elasticity estimated under the hypothesis of monopolistic competition and economies of scale has a significant and positive impact on trade, with a magnitude near to 5. This estimated coefficient is a short-run trade elasticity to the preferential margin factor. Thus, a one percentage point increase in the preferential margin factor is associated with a 5% increase in rice exports to the European Union, *ceteris paribus.* This result could be compared with the findings of Cipollina and Salvatici (2010), whose empirical estimation is based on agricultural commodity imports from 161 countries to the EU in 2004. Using the Heckman procedure but in a (cross-country) static setting, thus without taking into account issues of endogeneity or dynamics, they find trade elasticity to preferential margin equal to 3.8 for cereals and the cereal preparation sector.

The long-run elasticity can be obtained by dividing the coefficient by $(1-\gamma_1)$, where $\gamma_1$ is the coefficient of the lagged dependent variable. Thus, the consequent long-run effect of the preferential margin

[19] Standard results for omitted variable bias indicate that the OLS level estimator is biased upward, while the within estimator is biased downward (Bond, 2003). As a result, the consistent estimator should lie between OLS and within groups estimate.

[20] Only first order correlation is present, indicating inconsistency in the OLS estimator.

factor on trade is near 13, a magnitude that confirms the inertial behaviour of exports possibly due to sunk costs. We can now compare the long term *PM* elasticity obtained from the dynamic model with the results obtained using the static model. The estimated coefficient of $PM^E$ obtained from the LSDV static estimator (Column 3 of Table 4) is equal to 11, thus very close to the long-run elasticity obtained from the System-GMM.

Finally, Table 6 reports the results when three different preferential groups of countries are considered separately (ACP-OCT, EBA, Egypt).[21] In line with the results in Table 5, the preference impacts estimated using $PM^P$ are not significant, while those based on $PM^E$ are all significant, except one. In particular, the strongest impact of preferences on trade is found in ACP-OCT, with a magnitude of 10.4, followed by EBA countries. The positive and statistically significant coefficients indicate that the EU preferences matter, but there are some differences. Indeed, for ACP-OCT countries the cut in $PM^E$, after 2004 and, softened by the reduction of both MFN and preferential tariffs (see section 3), preserves a strong short-run impact on trade. On the contrary, for EBA countries, whose preferences drastically decreased after 2004, we detect a minor short-run average effect of preferences on trade flows. Finally, as regards Egypt, the *PM* coefficient is not statistically significant and this seems to indicate the lack of dependence of Egypt's rice exports on EU preferences.

Overall, from our estimates of the dynamic model using $PM^E$ , following the 2003 reform of the EU policy the preferential margin for the ACP countries declined in one year (2004-05) by 68.7% and this reduced ACP rice exports toward the EU by 18.8%, which is a rather considerable impact. As for EBA countries, the margin declined in one year by 76.6%; however, despite this sharp reduction in the margin their rice exports toward the EU fell only by 1.1% between 2004 and 2005.

## 7. Concluding remarks

Preference erosion is a key issue in the trade relationships between the EU and developing countries. Besides progress in multilateral liberalization under the WTO, there are other reasons why the preferences granted by the EU to developing countries are declining; the change in the EU trade policy in the rice industry in the past decade is one example. Although it is evident to many observers that the changes in EU trade policy have implied some erosion of preferences in the rice industry – one of the most sensitive industries for a number of developing countries – to date there has been no quantitative assessment of the extent of this erosion or on its impact on trade. This paper has addressed these issues, following two main strands: first, by proposing a new empirical approach to calculate the preferential margin when tariff rate quotas are in force; second, by assessing the trade impact of the preferential margins by means of both static

[21] To isolate the impact of preferential margin on the three groups we remove the export flows from the two preferential country groups other than the one analysed and maintain exports from non-preferential countries as the benchmark.

and dynamic panel gravity models to deal with endogeneity of the preferential margin and persistency in bilateral trade flows.

The results show that, when dealing with highly disaggregated data such as in this paper, the use of the “standard” tariff equivalent, i.e. the one consistent with the assumption of perfect competition and increasing marginal costs, may lead to an overestimation of the tariff, and thus to an underestimation of the preferences, when there are economies of scale in international trade. Further, on the basis of the value of the “standard” preferential margins one could conclude that no erosion of preferences has occurred, while this is not the case when using the preferential margins based on the tariff equivalent proposed in this paper. Thus, the main implication of this part of our analysis is that, when preferences are granted in the form of TRQs, the implicit assumptions on the market structure are very important. Although, to the best of our knowledge, no empirical evidence is available on the market structure and the cost curve of EU rice importers, we do believe that the existence of fixed costs and, thus, of the presence of economies of scale in the international trade of agricultural products are reasonable assumptions. In this case the use of the “standard” tariff equivalent of tariff rate quotas may result in misleading conclusions about the extent of the trade preferences and preference erosion.

The second major finding of this paper is that EU preferences matter significantly as regards the ability of developing countries to export rice to the EU. Trade elasticities are always lower with the standard tariff equivalent based on perfect-competition used in previous papers and, importantly, they are never statistically significant once dynamic is introduced into the model. These results suggest that the assumption on market structure is also important when assessing the trade impact of preferences. Finally, we find heterogeneity in trade preference elasticities across country groups, with ACP-OCT countries showing significantly higher values than the EBA group, while Egypt’s rice exports to the EU appear less dependent on preferences.

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Table 1. Preferential TRQs granted by the EU in the rice industry

| | Product | In-quota tariff | Quota |
|---|---|---|---|
| **ACP** | | | |
| *2000-2003* | Paddy and husked rice<br>Milled rice | $(T^{MFN}*0.5)$ - 4.34 Ecu/t<br>$(T^{MFN}$ - 16.78 Ecu/t)*0.5 - 6.52 Ecu/t | 125,000 t |
| | Broken rice | $(T^{MFN}*0.5)$ - 3.62 Ecu/t | 20,000 t |
| *2003-2008* | Paddy and husked rice<br>Milled rice | $(T^{MFN}*0.35)$ - 4.34 €/t<br>$(T^{MFN}$ - 16.78 Ecu/t)*0.35 - 6.52 €/t | 125,000 t |
| | Broken rice | $(T^{MFN}*0.35)$ - 3.62 €/t | 20,000 t |
| **OCT** | | | |
| *2000-08* | Paddy, husked and milled rice | 0 | 35,000 t |
| **Egypt** | | | |
| *2000-08* | All products | $T^{MFN}*0.75$ | 32,000 t |
| *2005-2008* | All products | 0 | 5,605 t |
| *2008* | Husked rice<br>Milled rice<br>Broken rice | 11 €/t<br>33 €/t<br>13 €/t | 57,600<br>196,000 t<br>5,000 t |
| **Bangladesh** | | | |
| *2000-2007* | Paddy and husked rice<br>Milled rice | $(T^{MFN}*0.5)$ - 3.6 Ecu/t<br>$(T^{MFN}*0.5)$ - 5.4 Ecu/t | 4,000 t |
| *2008* | Paddy and husked rice<br>Milled rice | $(T^{MFN}*0.5)$ - 4.34 €/t<br>$(T^{MFN}$ - 16.78 Ecu/t)*0.5 - 6.52 €/t | 4,000 t |
| **EBA** | | | |
| *2002-2008* | All products | 0 | from 2,895 t to 6,694 t |

*Source* : ACP and OCT: Regulations (EC) n. 2603/1997 and n.2286/2002. Egypt: Regulations (EC) n.196/1997, n.1002/2007, n.955/2005 and n. 1455/2007. Bangladesh: Regulations (EC) n. 3491/1990 and n.1532/2007. For EBA: Regulation (EC): 1401/2002

Table 2. EU imports of husked rice from Guyana: preferential margins under different hypotheses

| | Over quota imports (*ton*) | *Relative margin (%)* | | *Absolute margin (%)* | |
|---|---|---|---|---|---|
| | | Weighted - $PM^E$ | Standard-$PM^P$ | Weighted - $PM^E$ | Standard-$PM^P$ |
| 2000 | 0 | 18.1 | 18.1 | 21.7 | 21.7 |
| 2001 | 96 | 19.7 | 9.0 | 24.4 | 12.2 |
| 2002 | 0 | 24.7 | 24.7 | 32.1 | 32.1 |
| 2003 | 23551 | 23.3 | 0.0 | 31.2 | 0.0 |
| 2004 | 4741 | 25.1 | 0.0 | 29.1 | 0.0 |
| 2005 | 9733 | 7.8 | 0.0 | 8.1 | 0.0 |
| 2006 | 0 | 9.7 | 9.7 | 10.0 | 10.0 |
| 2007 | 2806 | 10.2 | 0.0 | 10.7 | 0.0 |
| 2008 | 0 | 7.1 | 3.4 | 7.3 | 3.6 |

*Source*: our computations based on Eurostat and EC Commission (2008). See text.

Table 3. The trade effect of preferential margin: Cross-section regressions

| | Dep. Variable: ln($import_{jkt}$) | | | | | |
|---|---|---|---|---|---|---|
| | $PM^P$ - Standard | | | $PM^E$ - Weighted | | |
| | 2001 | 2005 | 2008 | 2001 | 2005 | 2008 |
| **log(1+$PM_{jkt}$)** | -0.14 | -2.51 | 4.66 | 4.38 | 9.72 | 14.24** |
| | (7.15) | (4.73) | (6.59) | (6.38) | (7.74) | (6.07) |
| No. of obs. | 300 | 363 | 425 | 300 | 363 | 425 |
| R-Sq | 0.43 | 0.47 | 0.57 | 0.43 | 0.47 | 0.57 |

*Notes*: Exporter, and 6-digit product fixed effects included in each regression. Robust standard errors in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively.

Table 4. The trade effect of preferential margin: Static panel regression (2000-2008).

| *Dependent variable:* | ln($Import_{jkt}$) | | | | $Import_{jkt}$ | |
|---|---|---|---|---|---|---|
| | LSDV | | HECKMAN | | PPML | |
| | Standard-$PM^P$ | Weighted-$PM^E$ | Standard-$PM^P$ | Weighted-$PM^E$ | Standard-$PM^P$ | Weighted-$PM^E$ |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| **log(1+$PM_{jkt}$)** | 4.91** | 11.45*** | 20.54*** | 20.75*** | 10.64*** | 18.36*** |
| | (2.28) | (2.14) | (4.41) | (5.19) | (1.90) | (1.38) |
| Mills ratio | | | 3.37*** | 1.85** | | |
| | | | (0.74) | (0.82) | | |
| No. of obs. | 3.195 | 3.195 | 3.195 | 3.195 | 17.944 | 17.944 |

*Notes*: Exporter-year, time and 6-digit product-time fixed effects included in regressions (1)-(4). Exporter-year, time and 6-digit product fixed effects included in regressions (5)-(6) (see text). Robust standard errors clustered by country-pair in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively.

Table 5. Dynamic panel model: LSDV vs. system-GMM (2000-2008)

| | LSDV-dynamic | | Sys-GMM | |
|---|---|---|---|---|
| | Standard $PM^P$ | Weighted $PM^E$ | Standard $PM^P$ | Weighted $PM^E$ |
| | (1) | (2) | (3) | (4) |
| **log(trade$_{jk(t-1)}$)** | 0.79*** | 0.79*** | 0.61*** | 0.61*** |
| | (0.02) | (0.02) | (0.10) | (0.07) |
| **log(1+PM$_{jkt}$)** | 2.45 | 4.36** | 7.97 | 5.03** |
| | (1.75) | (1.72) | (7.64) | (2.13) |
| **log(distance$_j$)** | | | -0.23 | -0.17 |
| | | | (1.96) | (1.01) |
| **log(production$_{jt}$)** | | | 0.15 | 0.16** |
| | | | (0.11) | (0.06) |
| **constant** | 2.23*** | 2.23*** | 5.23 | 4.89 |
| | (0.29) | (0.29) | (16.54) | (8.17) |
| No. Obs. | 1910 | 1910 | 1683 | 1683 |
| No. Groups | | | 390 | 390 |
| No. Instruments | | | 83 | 83 |
| AR(2) | | | 0.273 | 0.264 |
| Hansen p-value | | | 0.764 | 0.709 |
| diff-in-Hansen p-value | | | 0.436 | 0.692 |

*Notes*: Time and 6-digit product-time fixed effects included in each regression. Robust standard errors in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively. We used all variables as instruments in model. The System-GMM estimator is implemented in STATA using the xtabond2 routine with option laglimits (3).

Table 6. Dynamic panel model: results across different preferential groups (2000-2008)

| | System- GMM | | | | | |
|---|---|---|---|---|---|---|
| | Standard $PM^P$ | | | Weighted $PM^E$ | | |
| | ACP-OCT | EBA | EGYPT | ACP-OCT | EBA | EGYPT |
| **log(trade$_{jk(t-1)}$)** | 0.70*** | 0.67*** | 0.65*** | 0.63*** | 0.67*** | 0.64*** |
| | (0.09) | (0.09) | (0.08) | (0.08) | (0.08) | (0.08) |
| **log(1+PM$_{jkt}$)** | 2.40 | 3.74 | 17.27 | 10.36* | 3.70** | -3.82 |
| | (7.84) | (10.86) | (27.94) | (5.49) | (1.48) | (51.89) |
| **log(distance$_j$)** | 0.13 | 0.16 | 0.16 | 0.11 | 0.13 | 0.07 |
| | (0.15) | (0.16) | (0.17) | (0.15) | (0.16) | (0.19) |
| **log(production$_{jt}$)** | 0.11*** | 0.13*** | 0.14*** | 0.16*** | 0.14*** | 0.14*** |
| | (0.04) | (0.04) | (0.04) | (0.05) | (0.04) | (0.04) |
| **constant** | 2.12 | 1.57 | 1.79 | 2.35* | 1.81 | 2.36 |
| | (1.47) | (1.52) | (1.53) | (1.41) | (1.53) | (2.10) |
| No. Obs. | 1501 | 1559 | 1517 | 1501 | 1559 | 1517 |
| No. Groups | 336 | 359 | 341 | 336 | 359 | 341 |
| No. Instruments | 82 | 72 | 75 | 84 | 77 | 80 |
| AR(2) | 0.302 | 0.399 | 0.320 | 0.311 | 0.371 | 0.309 |
| Hansen p-value | 0.860 | 0.697 | 0.757 | 0.949 | 0.847 | 0.881 |
| diff-in-Hansen p-value | 0.168 | 0.637 | 0.291 | 0.517 | 0.769 | 0.536 |

*Notes*: Time and 6-digit product-time fixed effects included in each regression. Robust standard errors in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively. The System-GMM estimator is implemented in STATA using the xtabond2 routine with option laglimits (3).

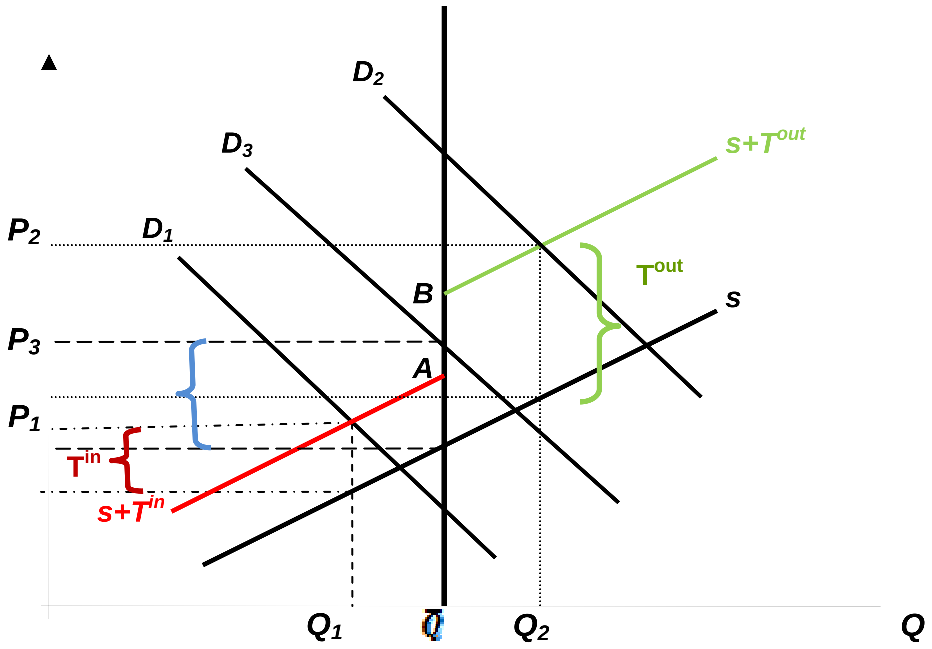


Figure 1: The tariff equivalent of a TRQ under perfect competition and increasing costs

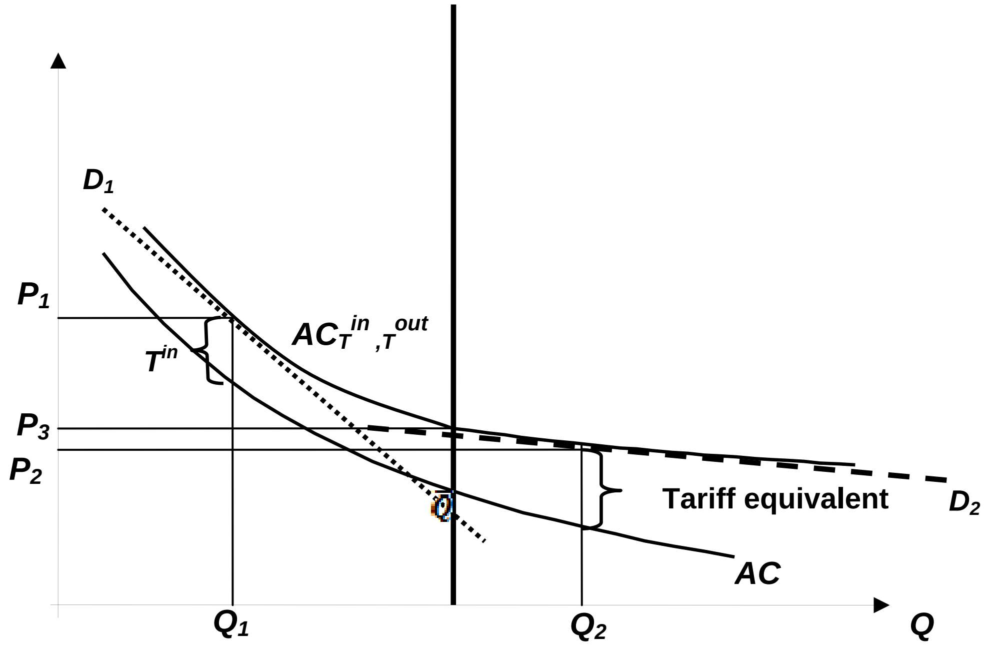


Figure 2: The tariff equivalent of a TRQ under economies of scale

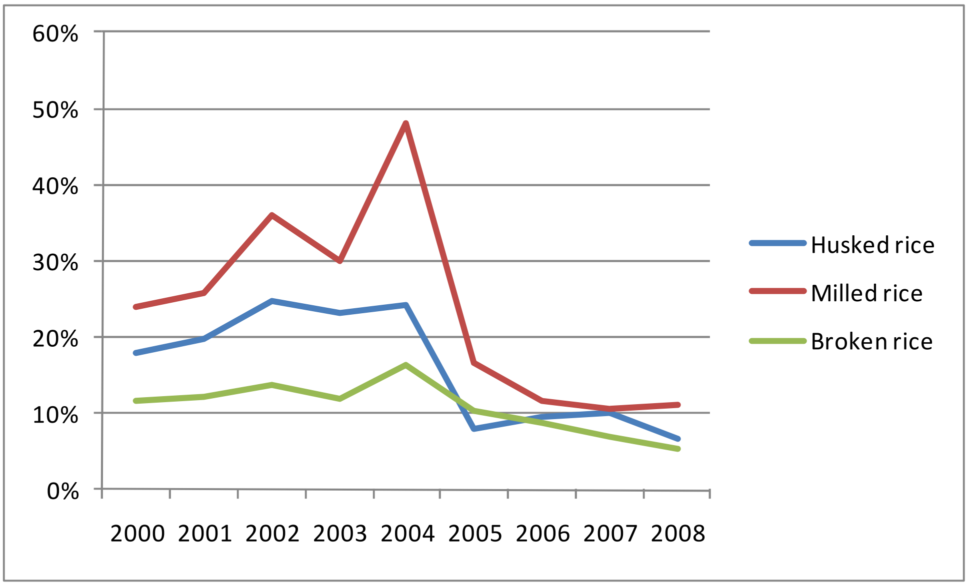


Figure 3: Average preferential margins under the assumption of economies of scale

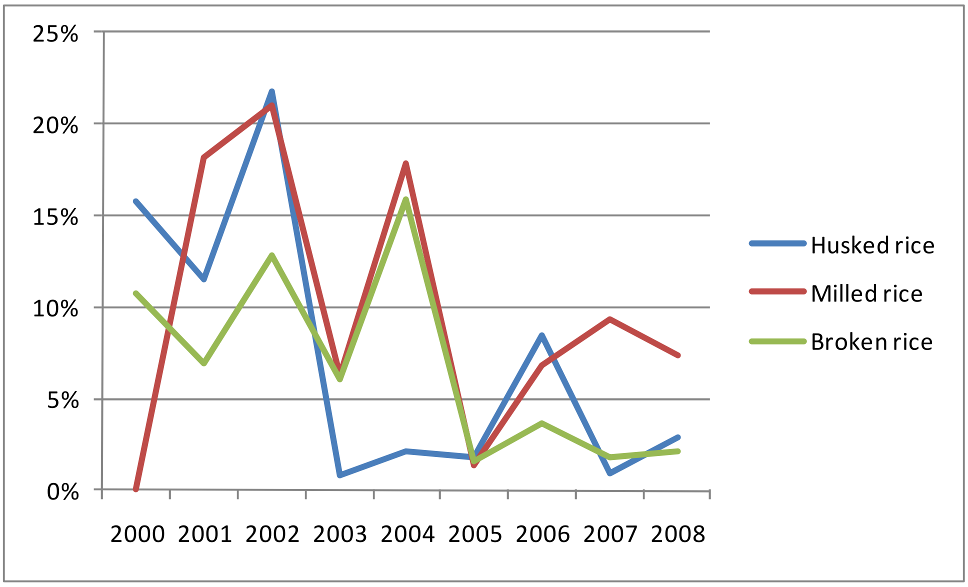


Figure 4: Average preferential margins under the assumption of perfect competition

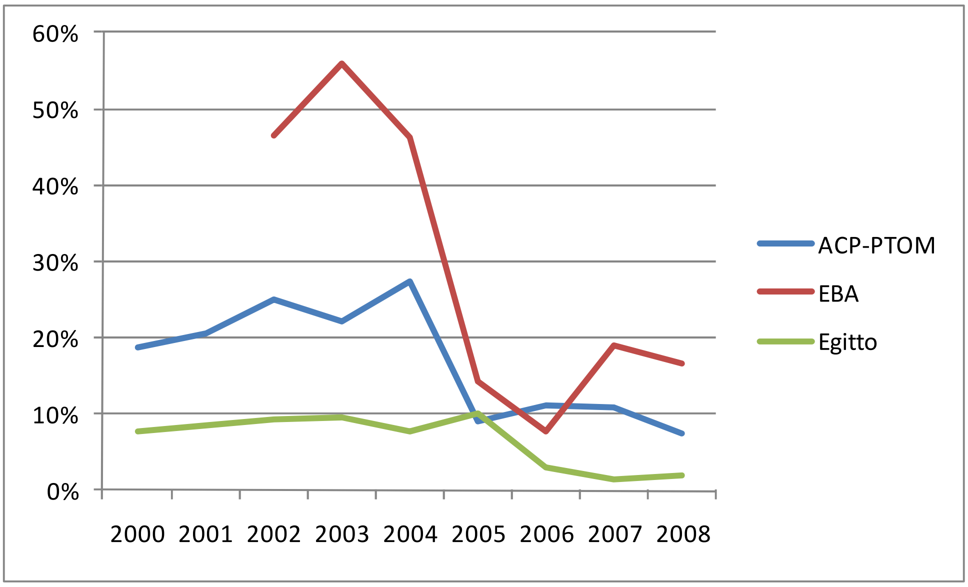


Figure 5: Average preferential margins by countries, under the assumption of economies of scale

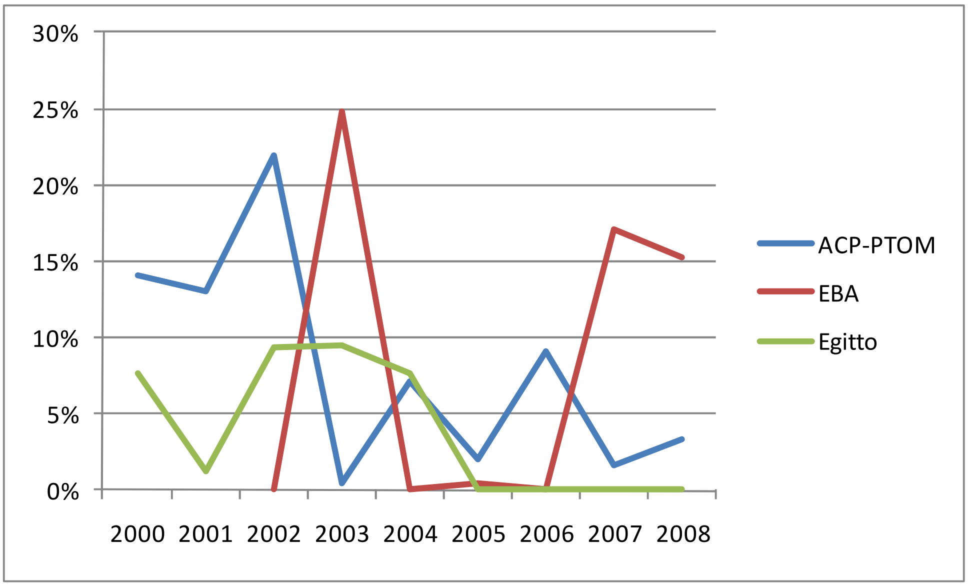


Figure 6: Average preferential margins by countries under the assumption of perfect competition