FOODSECURE
FOR POLICIES THAT MATTER

Are EU trade preferences really effective?

A Generalized Propensity Score evaluation of the Southern Mediterranean countries' case in agriculture and fishery

Emiliano Magrini
Pierluigi Montalbano
Silvia Nenci

FOODSECURE Working paper no. 23
June 2014

INTERDISCIPLINARY RESEARCH PROJECT
TO EXPLORE THE FUTURE OF GLOBAL
FOOD AND NUTRITION SECURITY

Are EU trade preferences really effective? A Generalized Propensity Score evaluation of the Southern Mediterranean Countries' case in agriculture and fishery

Emiliano Magrini[a], Pierluigi Montalbano[b,*], Silvia Nenci[c]

[a]*European Commission - JRC-IPTS - AGRILIFE Unit (Sevilla, ES)*
[b]*Department of Economic and Social Sciences - Sapienza, University of Rome (RM), Italy*
[c]*Department of Economics, University of Roma Tre, Rome (RM), Italy*

Abstract

The aim of this work is to assess the trade impact of preferential schemes. It focuses on the controversial case of the trade preferences in agriculture and fishery granted by the European Union (EU) to the Southern Mediterranean Countries (SMCs) over the period 2004-2009. The analysis presents several methodological improvements on previous works. Firstly, we rely on a continuous treatment - i.e. preferential margins - to capture the "average treatment effect" of trade preferences, rather than on a binary treatment based on dummy variables. Secondly, we apply a non-parametric matching technique for continuous treatment - specifically, a generalized propensity score matching (GPS) technique - to assess the average causal effects of preferences on trade flows. Thirdly, we use highly disaggregated data at sectoral level in order properly to evaluate the preferential treatment, which is conceived to be applied at the product level. The results show that the impact of the EU preferences in agriculture and fishery granted to SMCs is significant on SMCs trade flows and that is better evaluated by using impact evaluation techniques. Our findings raise important issues for policy-making by mitigating the claimed efficiency of the EU trade policy in the Mediterranean area.

Keywords: International trade, EU-MED integration, Preferential trade agreement, Impact evaluation, Matching econometrics.

JEL classification: C21,F10, F13, F15

Acknowledgements: We are grateful to all the participants in the following Seminars and Conferences at which previous versions of the paper have been presented: the ETSG Conference held in Leuven (BE), 13-15 September 2012; the Seminar held at the Department of Economics, University of Roma III, 6t March 2013. We thank Antonio Accenturo, Luca De Benedictis and Michael Gasiorek for their useful comments and suggestions on an earlier draft. The research leading to these results has received funding from the European Union's Seventh Framework programme FP7/2007-2011 under Grant Agreement nr. 290693 FOODSECURE - Exploring the Future of Global Food and Nutrition Security. Note that the views expressed are purely those of the authors and may not in any circumstances be regarded as stating an official position of the European Commission.

*Corresponding author. Department of Economics, University of Sussex, Falmer (UK) and Department of Economics and Social Sciences, Sapienza University of Rome (Italy). Email: p.montalbano@sussex.ac.uk. Tel. +39-06-49910739; Fax +39-06-49690326

April 26, 2014

1. Introduction

In the past two decades, the number of Preferential Trade Agreements (PTAs) has increased more than four-fold, to around 300 active agreements today. Nearly all countries are currently members of at least one PTA. Most PTAs now cover a wide range of issues besides tariffs, including services, investment, intellectual property protection, and competition policy (WTO, 2011).

Since the seminal work of Viner (1950) the effects of PTAs on international trade have been extensively studied. Standard gains from trade have been traditionally associated with the notion of trade creation, while the discriminatory nature of PTAs has been associated with trade diversion. Most of the literature agrees in finding positive effects of PTAs on trade flows among members (e.g. (Baier and Bergstrand, 2007, 2009; Magee, 2008)); some works highlight trade diversion effects on non-member countries (see (Trefler, 2004; Romalis, 2007; Carrére, 2006; Lee and Shin, 2006)); a few studies find inconclusive evidence (see, for example, (Calvo-Pardo *et al.*, 2009; Freund, 2010)). The most recent empirical evidence of the effects of PTAs on trade suggests more cautious optimism. Fugazza and Nicita (2013) highlighting the importance of the relative (vs. absolute) market access conditions, demonstrate that the current system of preferences is characterised by lower relative preferential margins for member countries and a higher deterioration of market access conditions for non-members. In the same vein, Cirera *et al.* (2011) suggest a very small impact of the EU Generalised System of Preferences on exports from developing countries, and a negligible or even negative one when they consider the scope for trade diversification. Finally, Hoekman and Nicita (2011) suggest that the effects of relative preferences for low-income countries are on average very small and preferences are more important on a country by country basis. The heterogeneous outcomes of the above empirical studies are normally linked to heterogeneity in samples, time periods, model specifications and PTA characteristics (Foster *et al.*, 2011). However, also within the context of the EU-MED partnership, the empirical evidence from the applied literature remains ambiguous (see (De Wulf and Maliszewska, 2009) and (Jarreau, 2011) for a survey).

Our aim is to provide conclusive evidence on the causal impact of trade preferences by focusing on the controversial case of the impact of trade preferences in agriculture and fishery granted by the European Union (EU) to the Southern Mediterranean Countries (SMCs) via the new generation of the EU-MED Association Agreements (AAs) in the time period 2004-2009 (i.e. from the end of the 12 year transitional period for the majority of AAs to the eve of the economic and political turmoil in the area). Differently from previous literature, we believe that focusing on a single specific preferential scheme is much more informative than relying on world averages and provides a more accurate and detailed analysis of the functional form of the empirical relationship, better controlling for unobservable counfounders. Moreover, our analysis presents several methodological improvements on previous empirical works. Firstly, we rely on a continuous treatment - i.e. preferential margins - to capture the "average treatment effect" of trade preferences, rather than on a binary treatment based on dummy variables. This provides clear advantages: the continuous treatment variable controls for heterogeneity in the depth and coverage of the EU preferential regime across products and countries, as well as for the actual rate of preference utilisation. To this end, we use highly disaggregated data by products. Secondly, we apply non-parametric matching techniques for continuous treatment to assess the average causal effects of preferences on trade flows, thereby addressing the issue of endogeneity between PTAs, trade flows and their determinants. More specifically, we apply the generalized version of the propensity score matching technique, i.e. the Generalized Propensity Score (GPS). This estimates treatment effects conditional on observable determinants of "treatment intensity". Non-parametric matching techniques help to isolate the treatment from any other event specific to the country pairs, and they also take into account the presence of non-linearities in the relationship among preferences, trade flows and the covariates (Baier and Bergstrand, 2009; Montalbano and Nenci, 2014). GPS, in particular, enables us to present a dose-response function and to illustrate how bilateral trade flows at the product level actually respond to changes in continuous treatment within the treatment group. In this case, the advantage is twofold: we are not compelled to undertake the difficult task of constructing a control group with similar characteristics - which is mandatory with the binary treatment matching techniques - and we address empirically the awkward issue of preferences' utilisation and the relative impact of trade preferences characterised by different intensities. The GPS method has been recently applied to various impact evaluation problems lacking experimental conditions: e.g. the impact of labour market programmes (Kluve, 2010; Kluve *et al.*, 2012), regional transfer schemes (Becker *et al.*, 2012), foreign direct investments (Du and Girma, 2009), and also the relationship between migration and trade (Egger *et al.*, 2012). To the best of our knowledge this is the first application of GPS to the assessment of the trade impact of the EU-MED preferential policy.

Our paper assesses the functional form of the relationship between the EU preferences and SMCs trade flows to the EU. The results show that the impact of the EU preferences in agriculture and fishery granted to SMCs is significant on SMCs trade flows and that is better evaluated by using impact evaluation techniques.

The paper is organised as follows: Section 2 briefly summarises the literature; Section 3 reports some stylised facts on the EU-MED partnership; Section 4 presents the GPS estimator; Section 5 describes the variables and data; Section 6 shows the empirical results; Section 7 concludes.

2. Literature review

Two main methodological approaches are commonly applied in order to measure the impact of preferences scheme on trade flows: computable general equilibrium (CGE) models to quantify the impact *ex ante*, and gravity models to measure it *ex post*.

Both static and dynamic effects have been considered in recent CGE studies on economic integration (see, among others, (Lee and van der Mensbrugghe, 2008; Bouet *et al.*, 2008). The static model evaluates the one-off, more immediate impact of the removal of trade barriers (Gilbert *et al.*, 2001; Urata and Kiyota, 2005). The dynamic model incorporates medium-term to long-term efficiency gains from resource reallocation and capital accumulation (Cheong, 2003; Francois and Wignaraja, 2008). Yet these models have been criticised because PTA results have been shown to be particularly sensitive to assumptions on the trade elasticity (Brown *et al.*, 1992; Ackerman and Gallagher, 2008), and a further limitation of CGE models is their poor economic interpretation of trade policy effects because of their structural complexity and data requirements (Panagariya and Duttagupta, 2001).

In the gravity approach, the effect of trade agreements is usually estimated by including dummy variables to control for the presence of policy factors and assess the extent to which PTA partners trade more than would be predicted by standard bilateral trade determinants. The "dummy strategy" is in fact the most workable solution. But it is unsatisfactory for a number of reasons: it implicitly assumes equal treatment and does not control for gradual implementation of the agreements; it does not control for specific country pair events contemporaneous with PTAs; it is unstable and loses significance the more that heterogeneity in the model is controlled for. This strand of the empirical literature typically focuses on either the EU or the USA preferences schemes (see, among others, (Acharya *et al.*, 2011; Agostino *et al.*, 2007; Caporale *et al.*, 2012; Collier and Venables, 2007; De Benedictis *et al.*, 2005; Di Rubbo and Canali, 2008; Frazer and Van Biesebroeck, 2010; Martínez-Zarzoso *et al.*, 2009; Péridy, 2005; Nilsson, 2007; Nilsson and Matsson, 2009)). Other studies apply a continuous variable by computing various measures of "preferential margins" guaranteed by a country to its partners (Aiello and Demaria, 2010; Cipollina and Salvatici, 2010; Francois *et al.*, 2006; Hoekman and Nicita, 2011; Kee *et al.*, 2009; Low *et al.*, 2009; Nilsson and Matsson, 2009; Montalbano and Nenci, 2014). Most of the works that employ a dummy variable conduct empirical analysis on aggregate data, while papers that employ continuous variables focus on disaggregated data.

Because a PTA is not an exogenous random variable but it is likely to be endogenously determined by and correlated with the country-pair trade flows and its determinants (Baier and Bergstrand, 2007; Egger *et al.*, 2008), several authors have recently addressed the endogeneity issue by relying on the impact evaluation methods and, in particular, using non-parametric matching techniques based on the benchmark between treatment and control groups. Persson (2001) has been one of the first to adopt this technique. In his work he uses non-parametric matching estimators to estimate the effects of a common currency on trade. He provides alternative estimates of the treatment effect more robust to selection and non-linearities than the linear regression strategy previously adopted. More recently, Egger *et al.* (2008) have applied a difference-in-difference analysis based on matching techniques to estimate the impact of endogenous new regional trade agreement membership on trade structure within the OECD economies, finding a strong effect of these agreements on intra-industry trade. The analysis of Baier and Bergstrand (2009) provides the first nonparametric empirical estimates using matching econometrics of the cross-sectionally long-run effects of free trade agreements on members' trade volume. They find a narrower range (across years) and more economically plausible values of the long-run effects of free trade agreements on members' trade than parametric ones in cross-section.[1] Millimet and Tchernis (2009) use propensity score-matching estimators to assess the environmental effects of General Agreement on Tariffs and Trade/World Trade Organization membership and the impact of adopting the euro on bilateral trade, finding respectively that the WTO is beneficial for environmental measures and that the euro has a positive effect on bilateral trade. Recently, Montalbano and Nenci (2014) have presented non-parametric matching estimates assessing the trade policy impact of the EU-MED free trade area; Cooke (2013) adopts a matching framework to estimate the impact of the AGOA policy of the USA on the exports of the beneficiary countries; Egger *et al.* (2012) use a semi-parametric approach to evaluation of the functional form of the relationship between migration (stocks) and trade (bilateral imports).

Our aim is to contribute to this strand of the empirical literature by applying a generalized version of the propensity score matching technique to estimate the impact of EU preferences regarding SMCs trade, conditional on observable determinants of "treatment intensity".

3. The EU-MED preferential trade policy: stylised facts

About twenty years since the launch of the Barcelona Process, SMCs are today fully involved in the EU-MED partnership, except for Syria (including the Palestinian Authority which has an Interim Euro-Mediterranean Association Agreement, see Table A.1 in the Appendix).

[1] Baier *et al.* (2013) confirm the above evidence on using also a gravity equation of both intensive and extensive margins and employing a long panel dataset with a large number of country pairs, product categories (4 digit SITC) and economic integration agreements. They carry out a set of robustness checks to potential country-selection, firm heterogeneity and reverse causality.

The objective of the network of bilateral AAs between EU member countries and SMCs is to provide for the gradual establishment of a Mediterranean Free Trade Area (FTA) in accordance with the rules of the World Trade Organization (WTO). It foresees the free movement of goods between the EU member countries and SMCs with the gradual removal of customs duties after a transitional period of twelve years following the entry into force of the AAs. As a result, since 1995, SMCs have recorded a dramatic decrease in Most Favored Nations (MFN) customs duties (below 18 percent for agricultural products and 5 percent for non-agricultural products) (Femise, 2011).

While all SMCs industrial goods are currently EU duty free, a quite new trade liberalization process is currently in place in the agricultural sector[2]. In all SMCs, agriculture plays a major role and is a key resource for the long term sustainability of economic development. By contrast, agriculture is a relatively small sector in the EU economy: it accounts for only 1.1 percent of GDP and 5.1 percent of employment (Tangermann and von Cramon-Taubadel, 2013). Nevertheless, the political weight of the agricultural sector is still remarkable, since it represents about 25 percent of the value of agricultural production in Spain, Italy, Greece, Portugal, Malta and Cyprus (E.C., 2003). Also testifying to the political importance of agriculture is the EU's relatively high level of protection for the agricultural sector, as well as its considerable heterogeneity. The EU protection is mainly characterized by: tariff-rate quotas, seasonal quotas and tariffs, threshold prices (Chevassus-Lozza *et al.*, 2005), and it varies markedly among different products (Jacquet *et al.*, 2007). Liberalisation between the EU and SMCs in the agricultural and food processing sectors was initially partial. However, a widening of the scope of the agreements to include greater liberalisation in the agricultural, agro-food and services sectors, as well as a reduction in non-tariff barriers (NTBs), is currently in progress (Jarreau, 2011). Although broad agricultural liberalisation still lags behind, some products already benefit from preferences granted within the framework of preferential agreements: more than 80 percent of agricultural products imported from the Mediterranean countries now enter the EU market either duty free or at reduced rates. Reciprocally, one third of EU exports of agricultural products benefits from preferential treatment in the Mediterranean countries. Furthermore, a system of Pan-Euro-Med cumulation of origin - also covering agricultural and fishery products - operates between the EU and SMCs. [3]

[2]Liberalization of trade in services and investment, including the right of establishment, is also part of the Association Agreements' key objectives, as well as the establishment of bilateral dispute settlement mechanisms for trade matters.

[3]In July 2011 the 42 members of the EU-MED partnership adopted the "PanEuroMed Protocol on cumulation of origin", which allows economic operators to cumulate processing made in different countries of the region and thus obtain preferential treatment. This Pan-Euro-Med cumulation is based on a network of preferential agreements that define the preferential tariffs, sometimes awarded for a limited volume (within quotas). The tariff reduction and hence the preferential margin enjoyed by the countries varies considerably within these quotas.

Notwithstanding the on-going liberalisation process, trade between EU and SMCs is still largely asymmetric. While 95 per cent of EU agricultural and fishery imports comes from outside SMCs, for the majority of SMCs the EU is the key destination market for their agricultural and fishery exports (more than 70 per cent of agricultural and fishery exports in the case of Algeria, Israel, and Morocco, more than 50 per cent for Tunisia). Moreover, as is apparent from Fig. A.1 in the Appendix, since the Barcelona Declaration, EU-MED trade relations have increased in absolute terms but worsened relatively to the other EU main trade partners[4]. Should we conclude that EU trade preferences are not effective? The question is not a trivial one. If we compare the EU preferential margins granted to our sample of products originating from SMCs before and after the entry into force of each respective EU-SMCs AA, apparent is not only a reduction in the mean level of protection but also a reduction in their heterogeneity minimum/maximum levels, as well as in their dispersion around the mean measured by standard deviation (see Figs. A.3 and A.4 in the Appendix). However, if we plot the mean value of exports by product before and after the entry into force of each respective EU-SMCs AA, a high degree of heterogeneity is still apparent (see Figs. A.5 to A.11 in the Appendix). While on average some products actually increase their exports to the EU market, other products reduce their mean export value towards the EU. The impression is that the impact of EU preferences is not unambiguous.

4. The GPS estimator

The GPS estimator - originally proposed by Hirano and Imbens (2004) and Imai and van Dyk (2004) - is a generalisation of the binary treatment propensity score. It is a non-parametric method used to correct for selection bias in a setting with a continuous treatment by comparing units that are similar in terms of their observable determinants of "treatment intensity" within the treatment group. Hence, it does not require control groups. It is based on the following assumptions: for each i there is a vector of covariates X_i, a "treatment" received, $T_i \in [t_0, t_1]$ and a potential outcome, $Y_i = Y_i(T_i)$. Following Hirano and Imbens (2004) we assume: Y_i, T_i and X_i are defined on a common probability space; T_i is continuously distributed with respect to a Lebesgue measure on τ; $Y_i = Y_i(T_i)$ is a well defined random variable. For each i we postulate the existence of a set of potential outcomes, $Y_i(t)$, for $t \in \tau$ where τ is the interval $[t_0; t_1]$ referred to as the unit-level dose-response function. We are interested in the average dose-response function across all observations i that illustrates the expected value of the outcome variable conditional on continuous treatment

[4] To highlight the different performances of SMCs and CEECs, in the figure we adopt the EU15 group instead of the EU27 one as in the empirical analysis.

as follows:

$$D(t) = E[Yi(t)] \tag{1}$$

In this exercise we use index $i = 1, ..., N$ to indicate the 6-digit products traded from SMCs to the EU27 area and assume the unit-level dose-response of potential outcomes in terms of EU bilateral imports, Y_{it} as a function of the treatment $t \in \tau$, where t is the product-level preferential margin granted by the EU. Following Hirano and Imbens (2004), we define GPS as:

$$R = r(t, X) \tag{2}$$

where R is the propensity score, i.e. the conditional probability of receiving a specific level of treatment given the covariates, and which is estimated via the following standard normal model:

$$\widehat{R}_i = \frac{1}{\sqrt{2\pi\widehat{\sigma}^2}} exp\left[-\frac{1}{2\widehat{\sigma}^2}(t_i - \widehat{\beta}_0 - X\widehat{\beta}_1)^2\right] \tag{3}$$

The main purpose of estimating GPS is to create covariate balancing. However, the validity of R as a measure of similarity or dissimilarity across product-level observations depends crucially on the validity of a set of assumptions which are standard in impact evaluation literature. The first assumption is the randomness of the treatment, i.e. the "unconfoundedness" or "ignorability of the treatment". This means that, conditional on observable characteristics, the treatment can be considered as random. Unconfoundedness is a key assumption in analysing trade preferences since countries sharing preferential agreements are unlikely to be randomly chosen (Baier and Bergstrand, 2007; Egger *et al.*, 2008). Imbens (2000) shows that if the treatment assignment is weakly unconfounded given the observed covariates, then the treatment assignment is weakly unconfounded given GPS. In other words, the GPS has the following property:

$$X \perp 1\left\{T = t\right\} | r(t, X) (4)$$

Another common assumption is the "overlap assumption", i.e., the need to maintain an adequate balance of observations between treatment and control groups. Using GPS we can easily get rid of it since we do not rely on control groups but instead work across GPS strata of various "treatment intensities" on a continuous distribution. Another assumption is the "unique treatment assumption" which is ensured in this case by the high degree of standardisation of the EU AAs. Last but not least we should take account of the "non-interference assumption", i.e., possible biases in the relationship between treatment and outcomes due to interfering events, such as the standard "trade diversion" effect in PTAs. Since our focus is on assessing the impact of the EU preferences on SMCs export flows, our analysis is not affected by the likely trade diversion of non-member countries to the EU. On the other hand, the small shares of SMCs exports on EU imports across products actually reduce also the relevance of trade diversion from SMCs (the SMCs product shares higher than 10% of the EU imports are only 4% of total observations, see Fig. A.2 in the Appendix). A robustness check that omits those observations in the empirical analysis is provided in section 6.[5]

GPS removes the bias associated with differences in covariates in three steps. In the first step, the GPS is estimated and its balancing property checked. If balancing holds, product-level flows within GPS strata can be considered as identical in terms of their observable characteristics, independently of their actual level of treatment.[6] Then, two additional steps are needed to eliminate the bias associated with differences on the covariates (see Hirano and Imbens (2004) for a proof). The first one is estimation of the conditional expectation of the outcome as a function of two scalar, the treatment level T and the GPS $R, \beta(t, r) = E[Y|T = t, R = r]$. The second and final step is to estimate the average dose-response function (DRF) of the outcome (product-level SMCs exports to the EU) averaging the conditional expectation over the GPS at any different level of EU product-level preferential margins, as follows:

$$D(t) = E[\beta(t, r(t, X))] \tag{5}$$

Furthermore, we can estimate the varying marginal effects of the treatment by estimating the treatment effect function, which is the first derivative of the corresponding dose-response function.

5. Variables and Data

In this exercise we use three different sets of data: the 6-digit product level preferential margins applied by the EU to the SMCs (i.e., the treatment, T_i); the observable characteristics which explain the probability of reaching a specific level of preferential margin (X_i); and the outcome in terms of export flows from SMCs to the EU at the 6-digit product level corresponding to the level of treatment received ($Y(t)$). Table A.2 in the Appendix reports a full description and the sources of the data applied in our empirical exercise.

[5] It is worth noting that in our case the presence of trade diversion among SMCs would eventually reduce the estimated causal impact of EU-MED AAs on SMCs trade flows to the EU.

[6] Note that as long as sufficient covariate balance is achieved, the exact procedure for estimating the GPS is of secondary importance (Kluve *et al.*, 2012).

Regarding the continuous variable for the actual product-level preferential margin granted to SMCs in the framework of the EU-MED AAs, we apply here the following measure of preferential margin (PM):

$$PM_{jit} = \frac{\sum_{v} T_{vit}^{EU} * imp_{vit}^{EU}}{\sum_{v} imp_{vit}^{EU}} - T_{jit}^{EU} \; with \; v \neq j \tag{6}$$

where T is the minimum tariff applied by the EU to imports of product i[7] and imp are the EU bilateral imports. i indexes the HS 6 digit categories; j indexes SMCs while v indexes the exporters competing with country j to access the EU market; t stands for observed years. While the second term (T_{jit}^{EU}) is simply the minimum tariff applied by the EU to imports of product i from country j, the first term is the counterfactual. The use of the counterfactual acknowledges the fact that for a given country it is the relative preference (i.e., the market access conditions relative to those faced by foreign competitors) that matters, not the absolute one, especially in the case of EU, because of the proliferation of EU PTAs around the world. It builds on the arguments put forward by Low *et al.* (2009); Carrére *et al.* (2010); Hoekman and Nicita (2011) and it is computed as the trade weighted minimum tariff level that the EU imposes on all other countries except j for which the preferential margin is calculated. Weights are the EU bilateral imports of product i from countries v, so as to take into account the supply capacity of SMCs competitors to the EU market. To mitigate the endogeneity problem we keep trade weights fixed over time in all the observed years (2004-2009) by taking the average values 1996-2003 of the EU bilateral imports. The PM_{jit} provides the relative advantage of SMC j in product i and year t with respect to each trading partner, capturing the discriminatory effects of the overall EU-MED system of preferences. Hence, it provides a reliable measure for the actual differences in EU market access both among SMCs - within the EU-MED AAs framework - and between SMCs and trade competitors eventually joining other PTAs with the EU. Moreover, the use of the applied tariffs controls directly for the actual utilisation rates of preferences, while the product level analysis controls for heterogeneous preferences on different products' origins. PM could be positive or negative depending on the advantage or disadvantage of the SMC_j in product i for each year t with respect to all the other competing exporters to the EU. It varies between the maximum negative bias (i.e. being the only trading partner facing tariffs when all other exporters enjoy duty free access) and the maximum positive bias (i.e., being the only trading partner enjoying duty free access while all other exporters face MFN tariffs). PM is zero when is no discrimination (i.e. the EU applies identical tariffs across all trading partners including duty free access). To be noted is

[7] This tariff rate is equal to the MFN applied tariff unless a preferential tariff exists in the database.

that the use of PM resolves a number of weaknesses of the simple dummy strategy. Firstly, it allows us to rely on a continuous measure of trade preferences. Secondly, it considers both the presence of differentiated treatments in the EU-MED framework and the issue of the gradual implementation of the EU-MED AAs. But it has drawbacks as well: it does not take into account the restrictive effects on non-tariff measures; and it takes into account only the direct price effects of tariffs, ignoring the general equilibrium of cross price effects (Fugazza and Nicita, 2013).

The issue of the covariates able to explain the probability of reaching a specific level of preferential margin is a controversial one. As stated by Baier and Bergstrand (2004), PTAs may well be a response to, rather than a source of, large trade flows, giving rise to endogeneity bias in trade impact evaluations. By introducing asymmetric absolute and relative factor endowments into a Krugman-type increasing-returns/monopolistic-competition model, they present, theoretically and empirically, the following determinants of the likelihood of bilateral PTAs: countries' economic size, distance, trade similarity and relative factor endowments.[8] Following the early literature on the sectoral determinants of trade protection after Finger (1981), Olarreaga and Vaillant (2011) explore the role played by microeconomic and macroeconomic variables in explaining the determinants of temporary trade barriers at the product level for Brazil. The price and value of imports at the product level are the main macroeconomic determinants that they use. Furthermore, to control for the role played by other microeconomic determinants pertaining to political economy (such as the concentration of the sector, output, or the extent to which workers are unionized), they use fixed effects and time varying effects. Among macroeconomic determinants they focus on MFN tariffs, real bilateral exchange rates (as a measure of competitiveness), and traditional measures such as the level of economic activity, unemployment, and institutional changes. Karacaovali and Limao (2008) emphasize the role of political economy, together with more standard economic variables, such as trade elasticity, market access, world price, and scale economies. Lastly, Gawande *et al.* (2011) underline the influence of both political economy variables (such as WTO-bound tariff, import demand elasticity, output-to-import ratio), and trade and product specialization measures (such as intra-industry trade, intermediate output, vertical specialization) in explaining the (weak) demand for protectionism.

In this empirical exercise we control for the actual determinants of preferential margins by assuming three main channels of impact. The first channel controls for country and product specific characteristics as well as time variant events in determining likely sources of protectionism. To this end, a full set of countries, time

[8] Baier and Bergstrand (2004) correctly predict, solely on the basis of economic characteristics, 85 percent of the 286 FTAs existing in 1996 among 1431 pairs of countries and 97 percent of the remaining 1145 pairs with no FTAs.

and products fixed effects as well as the HS6 digit EU import demand elasticities have been considered. To be noted is that in the period of observation there were no time-variant political and/or economic occurrences able to determine a shift in EU trade policy towards SMCs (both the political turmoil widely known as the "Arab spring" as well as the economic consequences of the recent financial crisis actually occurred after the period of analysis). Furthermore, in the same period the SMCs under analysis did not register any change in PTAs with other trade partners that could influence their trade flows with the EU.[9] The second channel controls for the role of *ad valorem* equivalents of quotas or other nontariff barriers (NTBs) seen as possible substitutes for preferences. Finally, since there is little scope for intra-industry and vertical specialisation in agricultural and fishery trade, we control for trade specialisation by relying on simple measures of SMCs trade specialisation, as well as EU import penetration by product. Moreover, as suggested by Baier and Bergstrand (2004), we also control for the GDPs difference between trading partners, since the probability of a PTAs is higher, the more similar they are.

As a measure of trade specialisation we apply the absolute and relative [10] product-level Lafay index as follows:

$$L_{jit} = \left[\frac{x_{it}^j - m_{it}^j}{x_{it}^j + m_{it}^j} - \frac{\sum\limits_i x_{it}^j - \sum\limits_i m_{it}^j}{\sum\limits_i x_{it}^j + \sum\limits_i m_{it}^j} \right] * \left[\frac{x_{it}^j + m_{it}^j}{\sum\limits_i x_{it}^j + \sum\limits_i m_{it}^j} \right] * 100 \tag{8}$$

This index measures country j's level of trade specialisation or revealed comparative advantage for each year t as the contribution to the trade balance of each product i to overall exports of country j. To acknowledge the meaning of the index, note that if there were no comparative advantage or disadvantage for any industry i, then country j's total trade balance (surplus or deficit) should be distributed across all industries according to their share in total trade. The 'contribution to the trade balance' is the difference between the actual and this theoretical balance. Hence, a positive contribution is interpreted as a 'revealed comparative advantage' for that industry. The advantage of using this index lies in its ability to derive a workable measure of each country's comparative advantages as they are revealed in trade data, avoiding difficulties linked to quantitative evaluations of factor-endowments and relative prices. Of course, decisions

[9] The only exceptions were Lebanon and Tunisia, which joined EFTA during the period of analysis. However, after controlling for this, the empirical results do not change.

[10] The relative version of the product-level Lafay index is:

$$L_{ijt} = \left[\frac{x_{it}^j - m_{it}^j}{x_{it}^j + m_{it}^j} - \frac{\sum\limits_j x_{it}^j - \sum\limits_j m_{it}^j}{\sum\limits_j x_{it}^j + \sum\limits_j m_{it}^j} \right] * \left[\frac{x_{it}^j + m_{it}^j}{\sum\limits_j x_{it}^j + \sum\limits_j m_{it}^j} \right] * 100 \tag{7}$$

about preferences are not driven exclusively by the revealed comparative advantages. There are a number of issues, mainly outside the field of economics, that can have a role in determining a specific preferences structure, also relative to other preferential schemes. To take these largely unobservable issues into account, here we also apply a set of country, year and product specific fixed effects to control for all unobserved determinants of PM.

Our outcome variable is the exports flow in agriculture and fisheries from SMCs to the EU disaggregated by export countries, products and year. In this exercise we use HS classification at the maximum disaggregation available (6 digit). Hence, we take into account agricultural, food and fishery products listed under chapters 1 to 24 of the Harmonized System Code (HS), Sections I-IV. To link products and tariffs properly, we use the WITS-TRAINS dataset. Since the aim of our empirical exercise is to examine whether the effect of the change in tariffs is stronger, the greater the advantage that it provides relative to other competitors, the choice of the time span (from 2004 to 2009) has been determined by the timetable of the transitional period of 12 years after the entry into force for most of the EU-MED AAs. The decision to limit the time period to 2009 also keeps the analysis out of the incidence of the both political and economic turmoil that has occurred in SMCs in the most recent years.

The complete dataset from 2004 to 2009 includes 1865 observations.[11] However, two more sample restrictions are applied. First, we eliminate those observations which can be considered as 'duty free access' cases (i.e. when both the counterfactual and the minimum tariff applied by the EU are equal to zero) or where corresponding data on covariates are not available. Second, we eliminate the observations in the first and the last 5 percentiles of the preferential margin distribution in order to clean our dataset of potential outliers. These two restrictions leave us with 1218 observations.

6. GPS estimation and results

6.1. GPS estimation and balancing property

The first step of our impact evaluation exercise is to estimate the GPS and test the "balancing property". The joint Jarque-Bera normality test does not reject the null hypothesis of normal distribution for the treatment variable.[12] Table 1 presents the outcomes of the first stage equation. The selected covariates prove to be important determinants of selection into treatment intensities across products and SMCs, including the

[11]It includes the available data on both trade flows and tariffs at 6-digit product level for our sample of countries.

[12]We apply here a zero-skewness Box-Cox transformation of the treatment variable. The p-value is equal to 0.11, above the 5% threshold of significance.

series of fixed effects previously described. The absolute Lafay index of the EU is positive and significant.[13] It indicates that the greater the contribution to the EU trade balance of a specific sector, the higher the probability that SMCs have a positive preferential margin within the EU's overall preferential scheme. Since here we are using a relative measure of preferential margin rather than the absolute one, it shows that, at a certain level of the protection of its products of relative specialisation, the EU prefers SMCs to their main competitors. Conversely, both the absolute Lafay index of the SMCs and the relative Lafay index between EU and SMCs are not significant.[14] This empirical evidence suggests that the EU-SMCs preferential scheme is mainly driven by the EU's trade policy strategy. Also the GDP coefficient is robust even if slightly positive.[15] The average EU import demand elasticity has a positive sign and it is highly significant. It shows the EU's relatively higher propensity to prefer SMCs for those products whose trade volumes are more sensitive to price variations. Finally, the positive sign and significance of the NTBs coefficient confirm the hypothesis of substitution between NTBs and tariffs. They empirically highlight the EU's strategy to maintain higher preferences in the sectors characterised by relatively higher quotas or other non-tariff measures.

Notwithstanding the relevance of our set of covariates, to be noted is that, in impact evaluation exercises, the interpretation and statistical significance of the individual effects of the covariates are of less importance than obtaining a powerful GPS (i.e., a GPS that works well in balancing the covariates by respecting the condition in eq. 4). In this regard, it is not irrelevant to add that the R-squared of our first stage regression is high and consistent with similar GPS empirical exercises (Becker *et al.*, 2012; Serrano-Domingo and Requena-Silvente, 2013).

Following the approach of Egger *et al.* (2012), we test the balancing property by comparing the covariates across groups with and without the GPS correction. Hence, we first perform a series of two-sided *t-tests* across groups for each covariate.[16] Groups of approximately the same size are formed on the basis of the actual preferential margin intensity, i.e. low (group 1); medium (group 2) and high (group 3). We obtain an average t-stat of 1.25. In 15.2% of the cases (119 over 783), the t-stat rejects the null hypothesis of equal mean among covariates, highlighting the presence of selection bias in the data. In order to remove this unbalance, as in Egger *et al.* (2012) and Becker *et al.* (2012), we exploit the results in the first stage

[13] Three lags of the Lafay index have been considered sufficient in this empirical exercise to avoid endogeneity problems.

[14] It is worth noting that removing the products where SMCs show higher revealed comparative advantages, the relative Lafay index becomes more robust, showing that the EU preferences are higher on the products where SMCs are less specialised (see Table A.4 in the Appendix).

[15] This result is consistent with Baier and Bergstrand (2004) because we use here the real GDPs of SMCs instead of the difference between them and that of the EU since in this case the EU GDP is a constant.

[16] This entails a remarkably high number of tests to be performed (261 variables for three groups, i.e. 783 t-tests), including fixed effects.

Table 1: **Generalised Propensity Score Estimation**

	coef	**SE (robust)**
EU Lafay Index	496.421***	204.83
SMC Lafay Index	1.265	3.31
EU-SMC Lafay Index	-0.030	0.035
GDP	0.001**	0.000
EU imports demand elasticity	0.131***	0.132
EU average non-tariff barriers	2.491**	1.179
Cons	1.623	2.684
SMCs Dummies	Yes	**15.82*****
Year Dummies	Yes	**3.19*****
6-digit product Dummies	Yes	**694.64*****
Observations	1 218	
R squared	0.49	

Note: In bold the results of the Wald tests for joint significance of the three groups of dummies. All time variant variables with three lags.

estimation (Table 1) to calculate the probability of each trade flow having the median preferential margin of the group T_M^j (for $j \in 1, 2, 3$), i.e. $\widehat{R}_i(T_M^j, X_i)$. We then plot these GPS values in group j against those not in group j (see figs. from A.12 to A.17) and eliminate those observations in groups other than j that lie outside the common GPS support. This means that we drop those trade flows in the control groups (blue bars) which lie outside the range of the treatment group (red bars), keeping only those flows which respect the following condition:

$$Min\widehat{R}_i(T_M^j, X_i) \leq \widehat{R}_l(T_M^j, X_l) \leq Max\widehat{R}_i(T_M^j, X_i) \quad (9)$$

for $i \in j$ and $l \notin j$.

The common GPS support condition is respected by 1085 observations out of 1218 used in the first stage, while 133 trade flows are pruned in order to ensure comparability among groups. We then organise the data in a group-strata structure to test the balancing property. This enables us to compare observations between treatment groups across strata based on the estimated GPS, i.e. to control for the ex-ante probability of receiving a specific preferential margin. For each of the three groups, six strata are determined on the basis of GPS scores evaluated at the median preferential margin of the respective group. We impose the same structure on the control observations in the same strata but in different groups. Table A.3 reports the final group-strata structure of the data.

To finalise the test of the balancing property, we again perform a series of *t-tests* comparing the mean difference in all the covariates among trade flows belonging to the same strata but in different groups (see Table A.5 in the Appendix). For example, we compare the trade flows belonging to Strata 1/Group 1 with the observations in Strata 1/Control 1. For each group, we then calculate the mean t-statistics weighting the t-stats by the number of trade flows in each stratum. After controlling for the GPS score, the average t-stat drops to 0.34 in the post-conditioning scenario. The number of cases where the t-stat rejects the null hypothesis of equal mean diminishes from 119 to 2. This means that our pre-treatment variables are well balanced among groups, confirming that the balancing property assumption holds and selection bias has been removed.[17]

6.2. Empirical results and robustness check

The last step of our empirical analysis is to estimate the DRF, i.e. to assess the level of SMCs' agriculture and fishery exports to the EU at any specific level of the observed preferential margin, given the estimated GPS. The GPS terms in this regression controls for selection into treatment intensities, while the interaction term shows the marginal impact of the treatment relative to the GPS. If selectivity matters, we expect both the GPS and the interaction coefficients to be statistically significant. This means that the GPS method highlights possible bias in outcomes that are actually controlled by looking over GPS strata, as well as across GPS by using the interaction term.

A number of polynomials can be tested to assess the above relationship. As in Egger *et al.* (2012) we chose to disregard polynomial terms that turned out to be insignificant.[18] The corresponding results for the parsimonious, semi-parametric DRF are summarised in Table 2. As emphasised by Hirano and Imbens (2004), similarly to what occurred in the first stage case, the parameters reported in Table 2 do not have a causal interpretation. Also in this case R-squared is relatively high, given the parsimonious specification, and consistent with similar GPS empirical exercises.

The key findings of Table 2 is that selection into treatment intensities is relevant in the EU-MED case. The GPS coefficient is positive and significant and the marginal impact of treatment intensity decreases along with GPS intensities, as shown by the negative sign of the interaction term coefficient.

The upper panel of Fig.1 reports the graphical representation of the point estimates of the DRF while

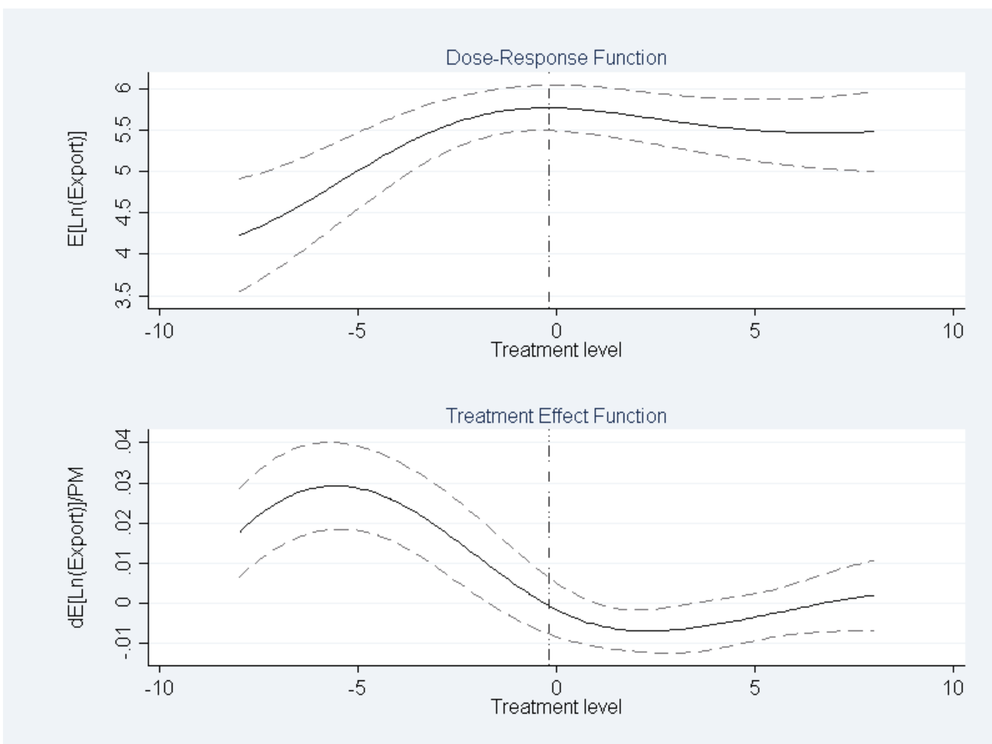
[17] We tested the balancing property for different group/strata structures. The results confirm that the preferred structure performs best in balancing the data.

[18] Other polynomial specifications neither add any relevant information nor affect the DRF.

Table 2: **Dose-Response Function estimation**

	Coef	**SE(robust)**
PM	0.190***	0.066
PM^2	0.024***	0.002
PM^3	-0.001***	0.000
GPS	303.265***	32.571
GPS^2	-5145.429***	1100.752
GPS^3	29235.28***	9113.825
$GPS * PM$	-6.024***	0.972
Observations	1 047	
R-squared	0.80	

Figure 1: The dose-response and the treatment effect function



the lower panel of Fig. 1 represents the treatment effect function (TRF), i.e. the first derivative of the DRF. The dashed lines are the corresponding 90% confidence intervals based on estimated bootstrapped standard errors. The upper panel of the figure shows the non linear relationship between the product level preferential margin granted by the EU in agricultural and fishery products and the respective level of SMCs exports to the EU. It is positive and increasing until it reaches the optimal level, which corresponds to neutrality. Then, positive EU preferences do not have an impact on SMCs' exports until they reach a very high level (i.e., over 5). This is the key features of our empirical analysis and the most important issue for policy-making. Consistently, the TRF in the lower panel of Fig.1 shows that the marginal change of SMCs exports in correspondence to a marginal change of the EU's preferences increases when the level of preferences is heavily negative, then decreases towards neutrality and increases again only from very high levels of EU preferences.

Table 3: **Dose-Response Function estimation (product shares on the EU imports less than 10%)**

	Coef	**SE(robust)**
PM	0.136**	0.068
PM^2	0.022***	0.002
PM^3	-0.000***	0.000
GPS	312.925***	31.336
GPS^2	-5458.23***	1047.38
GPS^3	30885.39***	8574.10
$GPS * PM$	-5.143**	1.011
Observations	983	
R-squared	0.80	

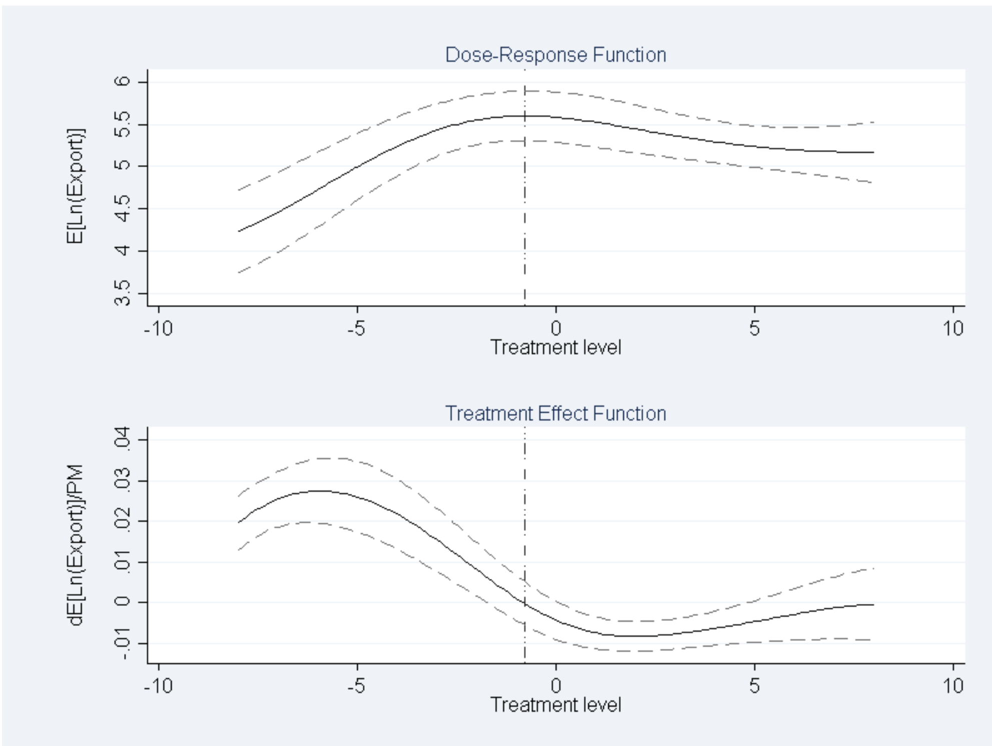
To avoid potential bias that may derive from trade diversion among SMCs - i.e., lower estimated impacts of preferences caused by diverted trade flows from the most relevant SMCs competitors - we perform the same exercise by dropping the SMCs products whose shares on EU imports are higher than 10% (which are supposed to be more affected by trade diversion). As is evident from Table 3 and Fig.2 there is no material difference in the empirical outcomes.[19]. The only noticeable difference between the two estimations is that in the latter one the optimal treatment is even lower with respect to neutrality.

7. Conclusions

The most recent debate on PTAs has focused on the following research question: do preferences impact on trade? While the common perception is that preferences do impact positively on trade, the empirical

[19]Due to space constraints the first stage outcomes are presented in Table A.4 in the Appendix

Figure 2: The dose-response and the treatment effect function (product shares on the EU imports less than 10%)



evidence is debated. The issue is becoming controversial within the framework of the EU-MED PTAs, since trade relations between EU and SMCs have worsened relatively to the EU's other main trade partners since the Barcelona Declaration. The aim of this work has been to assess the trade impact of EU-MED preferential schemes in agriculture and fishery by adopting a novel methodological approach, namely a GPS matching technique. Differently from the majority of current analyses, we chose a continuous variable to measure preferences in order to capture the "average treatment effect" of PTAs. Second, we applied non-parametric matching techniques for continous treatment to assess the average causal effects of preferences on trade flows. Third, we used highly disaggregated data at sectoral level in order properly to evaluate the preferential treatment which is conceived to be applied at the product level. Our paper has assessed the functional form of the relationship between EU-SMCs preferences in agriculture and fishery products and bilateral trade flows with continuous treatment under the (weak) unconfoundedness assumption. Our empirical results show that the impact of the EU product-level preferential policy on SMC trade flows in agricultural and fishery products is significant, and that is better evaluated by using impact evaluation techniques. They also suggest that while the removal of a relative disadvantage in terms of trade preferences has, on average and *ceteris paribus*, a positive effect on trade flows, positive preferences do not have a positive impact on trade until they reach a very high level (over 5). These findings raise important issues for policy-making mitigating the claimed efficiency of the EU trade policy in the Mediterranean area.

References

Acharya R., Crawford J.A., Maliszewska M. and Renard C. (2011) Landscape, in: *Preferential Trade Agreement Policies for Development: a Handbook*, Chaffour J.P. and Maur J., eds., World Bank, Washington DC.

Ackerman F. and Gallagher K. (2008) The Shrinking Gains from Global Trade Liberalization in Computable General Equilibrium Models: A Critical Assessment, *International Journal of Political Economy*, 37, 1.

Agostino M., Aiello F. and Cardamone P. (2007) Analysing the Impact of Trade Preferences in Gravity Models: Does Aggregation Matter?, *Trade AG working paper*, 07/4.

Aiello F. and Demaria F. (2010) Do Trade Preferential Agreements Enhance the Exports Of Developing Countries? Evidence from the EU GSP, *Working Papers, Universitá della Calabria, Dipartimento di Economia e Statistica*, 201002.

Baier S. L. and Bergstrand J. H. (2004) Economic Determinants of Free Trade Agreements, *Journal of International Economics*, 64, 29–63.

Baier S. L. and Bergstrand J. H. (2007) Do Free Trade Agreement Actually Increase Members' International Trade?, *Journal of International Economics*, 71, 72–95.

Baier S.L. and Bergstrand J.H. (2009) Estimating the Effects of Free Trade Agreements on International Trade Flows using Matching Econometrics, *Journal of International Economics*, 77, 63–76.

Baier S.L., Bergstrand J.H. and Feng M. (2013) Economic Integration Agreements and the Margins of International Trade, *University of Notre Dame, mimeo*.

Becker S.O., Egger P. and von Ehrlich M. (2012) Too Much of a Good Thing? On the Growth Effects of the EU's Regional Policy, *European Economic Review*, 56, 648–668.

Bouet A., Decreux Y., Fontagné L., Jean S. and Laborde D. (2008) Assessing Applied Protection across the World, *Review of International Economics*, 16, 5, 850–863.

Brown D., Deardorff A. and Stern R. (1992) A North American Free Trade Agreement: Analytical Issues and a Computational Assessment, *World Economy*, 15, 1, 11–29.

Calvo-Pardo H., Freund C. and Ornelas E. (2009) The ASEAN Free Trade Agreement: Impact on Trade Flows and External Trade Barriers, in: *Costs and Benefits of Regional Economic Integration*, Barro R. and Lee J.W., eds., Oxford University Press.

Caporale G.M., Rault C., Sova R. and Sova A. (2012) European Free Trade Agreements and Trade Balance: Evidence from Four New European Union Members, *The Journal of International Trade Economic Development*, 21, 6, 839–863.

Carrére C. (2006) Revisiting the Effects of Regional Trade Agreements on Trade Flows with Proper Specification of the Gravity Model, *European Economic Review*, 50, 2, 223–247.

Carrére C., de Melo J. and Tumurchudur B. (2010) Disentangling Market Access Effects of Preferential Trading Arrangements with an Application for ASEAN Members under an ASEAN-EU FTA, *The World Economy*, 33, 1, 42–59.

Cheong I. (2003) Regionalism and Free Trade Agreements in Asia, *Asian Economic Papers*, 2, 2, 145–180.

Chevassus-Lozza E., Gallezot J., Harel M. and Persillet V. (2005) The Protection of the European Market in the Fruit and Vegetable Sector. The Agreements between EU and Mediterranean Countries, *Eu-Med AgPol*, 1.

Cipollina M. and Salvatici L. (2010) The Trade Impact of European Union Agricultural Preferences, *Journal of Economic Policy Reform*, 13, 1, 87–106.

Cirera D.X., Foliano F. and Gasiorek M. (2011) The Impact of GSP Preferences on Developing Countries' Exports in the European Union: Bilateral Gravity Modelling at the Product Level, *Sussex Economics Department Working Paper Series*, 27.

Collier P. and Venables A. (2007) Rethinking Trade Preferences: How Africa Can Diversity its Exports, *The World Economy*, 30, 8, 1326–1345.

Cooke E.F.A. (2013) A Skeptics View of the AGOA Preferences of the USA: A Propensity Score Matching Approach, *MPRA Paper*, 39144.

De Benedictis L., De Santis R. and Vicarelli C. (2005) Hub-and-Spoke or else? Free Trade Agreements in the Enlarged EU, *European Journal of Comparative Economics*, 2, 245–260.

De Wulf L. and Maliszewska M. (2009) *Economic integration in the euro-mediterranean region*, (Eds.) Final Report, Center for Social and Economic Research and Center for European Policy Studies, September, Warsaw.

Di Rubbo P. and Canali G. (2008) A Comparative Study of EU and US Trade Policies for Developing Countries: the Case of Agri-food Products, *12th Congress of the European Association of Agricultural Economists*.

Du J. and Girma S. (2009) The Effects of Foreign Acquisition on Domestic and Export Markets Dynamics in China, *The World Economy*, 32, 1, 164–177.

E.C. (2003) *The Horticulture Sector in the European Union*, European Commission, DG Agriculture.

Egger H., Egger P. and Greenaway D. (2008) The Trade Structure Effects of Endogenous Regional Trade Agreements, *Journal of International Economics*, 74, 278–298.

Egger P.H., von Ehrlich M. and Nelson D.R. (2012) Migration and Trade, *The World Economy*, 35, 2, 216â241.

Femise (2011) *Report on the Euro-Mediterranean Partnership*, Femise, Marseille.

Finger J. M. (1981) The Industry-Country Incidence of Less-than-Fair-Value Cases in US Import Trade, *Quarterly Review of Economics and Business*, 21, 2, 260–279.

Foster N., Poeschl J. and Stehrer R. (2011) The Impact of Preferential Trade Agreements on the Margins of International Trade, *Economic Systems*, 35, 2, 84â97.

Francois J. F. and Wignaraja G. (2008) Economic Implications of Asian Integration, *Global Economy Journal*, 8, 3, 146.

Francois J., Hoekman B. and Manchin M. (2006) Preference Erosion and Multilateral Trade Liberalization, *The World Bank Economic Review*, 20, 197–216.

Frazer G. and Van Biesebroeck J. (2010) Trade Growth under the African Growth and Opportunity Act, *Review of Economics and Statistics*, 92, 1, 128–144.

Freund C. (2010) Third-country Effects of Regional Trade Agreements, *The World Economy*, 33, 11, 1589â1605.

Fugazza M. and Nicita A. (2013) The direct and relative effects of preferential market access, *Journal of International Economics*, 89, 2, 357–368.

Gawande K., Hoekman B. and Cui Y. (2011) Determinants of Trade Policy Responses to the 2008 Financial Crisis, *World Bank Policy Research Working Paper*, 5862.

Gilbert J., Scollay R. and Bora B. (2001) New Regional Trading Developments in the Asia-Pacific Region, in: *Global Change and East Asian Policy Initiatives*, Shahid Y., Altaf M.A. and Nabeshima K., eds., World Bank, Washington, DC.

Hirano K. and Imbens G.W. (2004) The Propensity Score with Continuous Treatments, in: *Applied Bayesian Modeling and Causal Inference from Incomplete-Data Perspectives: An Essential Journey with Donald Rubin's Statistical Family*, Wiley Series in Probability and Statistics, 73–84.

Hoekman B. and Nicita A. (2011) Trade Policy, Trade Costs, and Developing Country Trade, *World Development*, 39, 12, 2069–2079.

Imai K. and van Dyk D.A. (2004) Casual Inference with General Treatment Regimes: Generalizing the Propensity Score, *Journal of the American Statistical Association*, 99, 854–866.

Imbens G.W. (2000) The Role of the Propensity Score in Estimating Dose-Response Functions, *Biometrika*, 87, 3, 706–710.

Jacquet F., Petit M., Tyner W. and El Hadad F. (2007) *Impacts of Agricultural Trade Liberalization between the EU and Mediterranean Countries. EUMEDAGPOL Synthesis Report*, European Commission.

Jarreau J. (2011) Economic Integration in the EuroMed: Current Status and Review of Studies, *CEPII, WP*, 07.

Karacaovali B. and Limao N. (2008) The Clash of Liberalizations: Preferential vs. Multilateral Trade Liberalization in the European Union, *Journal of International Economics*, 74, 299–327.

Kee H.L., Nicita A. and Olarreaga M. (2009) Estimating Trade Restrictiveness Indices, *The Economic Journal*, 119, 172–199.

Kluve J. (2010) The Effectiveness of European Active Labor Market Programs, *Labour Economics*, 17, 6, 904–901.

Kluve J., Schneider H., Uhlendorff A. and Zhao Z. (2012) Evaluating Continuous Training Programmes by using the Generalized Propensity Score, *Journal of the Royal Statistical Society: Series A (Statistics in Society)*, 175, 2, 587–617.

Lee J.W. and Shin K. (2006) Does Regionalism Lead to More Global Trade Integration in East Asia?, *The North American Journal of Economics and Finance*, 17, 3, 283–301.

Lee J.W. and van der Mensbrugghe D. (2008) Regional Integration, Sectoral Adjustments and Natural Groupings in EastAsia, *International Journal of Applied Economics*, 5, 2, 57–79.

Low P., Piermartini R. and Richtering J. (2009) Multilateral Solutions to the Erosion of Non-reciprocal Preferences in NAMA, in: *Trade Preference Erosion Measurement and Policy Response*, Bernard Hoekman W.M. and Braga C., eds., Palgrave MacMillan.

Magee C. (2008) New Measures of Trade Creation and Trade Diversion, *Journal of International Economics*, 75, 2, 340–362.

Martínez-Zarzoso I., Nowak-Lehmann D.F. and Horsewood N. (2009) Are Regional Trading Agreements Beneficial? Static and Dynamic Panel Gravity Models, *North American Journal of Economics and Finance*, 20, 1, 46–65.

Millimet D. and Tchernis R. (2009) On the Specification of Propensity Scores: with Applications to the Analysis of Trade Policies, *Journal of Business and Economic Statistics*, 27, 3, 397–415.

Montalbano P. and Nenci S. (2014) Assessing the Trade Impact of the European Neighbourhood Policy on the EU-MED Free Trade Area, *Applied Economics*, 46, 7.

Nilsson L. (2007) Comparative Effects of EU and US Trade Policies on Developing Country Exports, *European Commission, DG Trade*.

Nilsson L. and Matsson N. (2009) Truths and Myths about the Openness of EU Trade Policies and the Use of EU Preferences, *European Commission, DG Trade*.

Olarreaga M. and Vaillant M. (2011) Micro and Macro Determinants of Trade Temporary Barriers: the Brazilian Case over the Last Two Decades, *Departemento de Economia, Universidad de la Repubblica, Working Paper*, 07.

Panagariya A. and Duttagupta R. (2001) The Gains from Preferential Trade Liberalization in the CGE Models: Where Do They Come from, in: *Regionalism and Globalization: Theory and Practice*, Lahiri S., ed., Routledge, London and New York.

Péridy N. (2005) The Trade Effects of the Euro Mediterranean Partnership: what are the Lessons for ASEAN Countries?, *Journal of Asian Economics*, 16, 125–139.

Persson T. (2001) Currency Unions and Trade: How Large is the Treatment Effect?, *Economic Policy*, 16, 433–462.

Romalis J. (2007) NAFTA's and CUSFTA's Impact on International Trade, *Review of Economics and Statistics*, 89.

Serrano-Domingo G. and Requena-Silvente F. (2013) Re-examining the Migration-trade Link using Province Data: An Application of the Generalized Propensity Score, *Economic Modelling*, 32, 247–261.

Tangermann S. and von Cramon-Taubadel S. (2013) Agricultural policy in the European Union: An overview, *Department f ur Agrar okonomie und Rurale Entwicklung Working Paper, Universit at G ottingen*, 1302.

Trefler D. (2004) The Long and Short of the Canada-U.S. Free Trade Agreement, *American Economic Review*, 94, 31.

Urata S. and Kiyota K. (2005) The Impacts of an East Asia Free Trade Agreement on Foreign Trade in East Asia, in: *International Trade in East Asia*, in Takatoshi Ito C.. and Rose A.K., eds., NBER East Asia Seminar on Economics.

Viner J. (1950) *The Customs Union Issue*, Carnegie Endowment for International Peace.

WTO (2011) *The WTO and Preferential Trade Agreements: from Co-existence to Coherence*, Geneva.

Appendix A.

Table A.1: **State of implementation of EU-MED AAs**

Country	Signature date	Entry into force
Algeria	22 April 2002	1 September 2005
Egypt	25 June 2001	1 June 2004
Israel	20 November 1995	1 June 2000
Jordan	24 November 1997	1 May 2002
Lebanon	17 June 2002	1 April 2006
Morocco	26 February 1996	1 March 2000
Palestinian Authority	24 February 1997	1 July 1997 (Interim association agreement)
Syria	Negotiations concluded awaiting for signature	
Tunisia	17 July 1995	1 March 1998
Turkey	6 March 1995	31 December 1995

Table A.2: **Variables and data sources**

Type	**Variable**	**Source**
Outcome variable	Export flows in agriculture and fisheries (Chapters 1 to 24 of the HS code, Sections I-IV)	BACII-CEPII
Treatment variable	Preferential margin (the 6-digit product level applied tariff)	WITS-TRAINS
Covariates:		
Institutions/political economy	Countries, time & products fixed effects; EU Import demand elasticities (HS6)	Kee, Nicita & Olarreaga, RES 2008; 2009
Trade specialisation	Absolute and relative Lafay index; GDPs difference	UN-COMTRADE; Penn World Tables 6.1
NTBs	Ad valorem equivalents of quotas or other nontariff barriers	Kee, Nicita & Olarreaga, RES 2008; 2009

Table A.3: **The final group-strata structure**

Strata	**Control1**	**Group1**	**Control2**	**Group2**	**Control3**	**Group3**
1	382	56	276	69	473	57
2	166	56	116	68	127	57
3	97	55	84	69	66	56
4	52	56	80	68	42	57
5	34	56	74	69	23	57
6	20	55	44	68	14	56
Total	751	334	674	411	745	340

Table A.4: **Generalised Propensity Score Estimation (product shares on the EU imports less than 10%)**

	coef	**SE (robust)**
L.EU Lafay Index	482.131**	190.494
L.SMC Lafay Index	0.649	3.677
L.EU-SMC Lafay Index	-0.042	0.052
LnGDP	0.002***	0.000
EU imports demand elasticity	0.045***	0.013
EU average non-tariff barriers	-1.232**	0.547
Cons.	-5.674**	2.737
SMCs Dummies	Yes	**14.73*****
Year Dummies	Yes	**4.01*****
6-digit product Dummies	Yes	**771.11*****
Observations	1 114	
R squared	0.50	

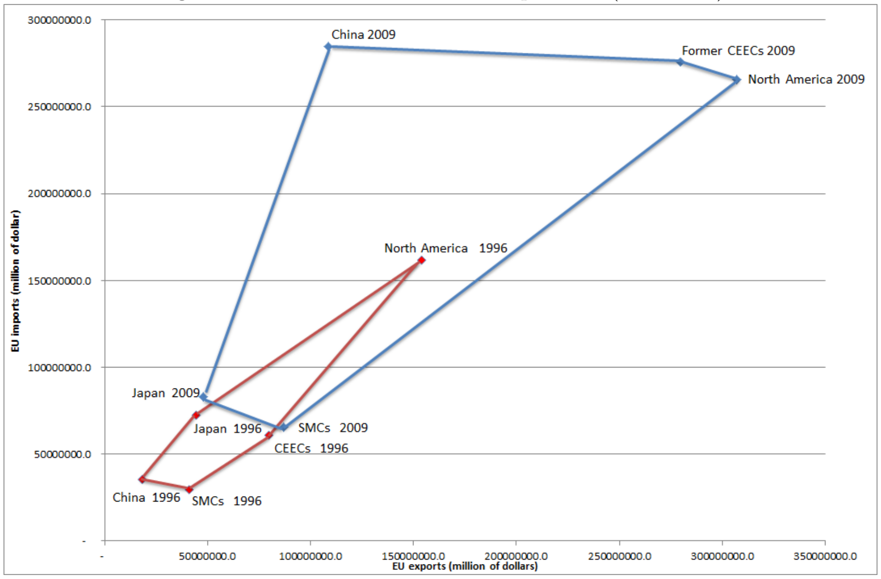
Note: In bold the results of the Wald tests for joint significance of the three groups of dummies. All time variant variables with three lags.

Table A.5: **Differences in treatment levels before and after balancing on the GPS: t-stats for equality of means.**

	Prior to balancing on the GPS			**After balancing on the GPS**		
Covariates	**Group 1**	**Group 2**	**Group 3**	**Group 1**	**Group 2**	**Group 3**
EU imports demand elasticity	1.313	**2.755**	1.458	0.603	0.623	0.228
EU average non-tariff barriers	1.577	1.308	**2.905**	0.268	0.714	0.410
EU Lafay Index	**2.874**	1.565	**4.479**	0.475	0.201	1.639
SMC Lafay Index	**2.144**	0.599	1.539	1.009	0.473	0.240
EU-SMC Lafay	1.108	0.085	1.194	0.341	0.015	0.529
Real GDP	**7.418**	**3.675**	**3.593**	0.070	0.458	0.338
Mean t-value		1.250			0.337	
Nr. unbalanced obs		119			2	

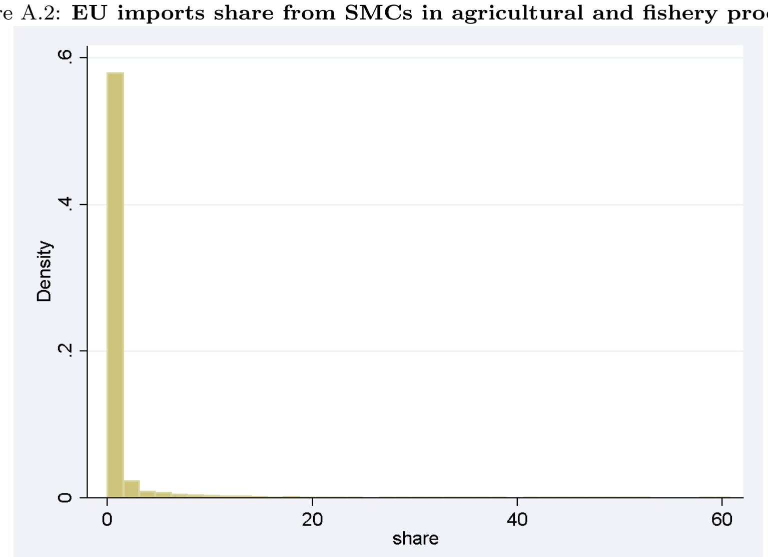
Note: t-values reported in bold face indicate null rejections at the 5% level significance.

Figure A.1: EU Trade with its main partners (1996-2009)



Source: Authors' own calculations on Comtrade

Figure A.2: **EU imports share from SMCs in agricultural and fishery products**

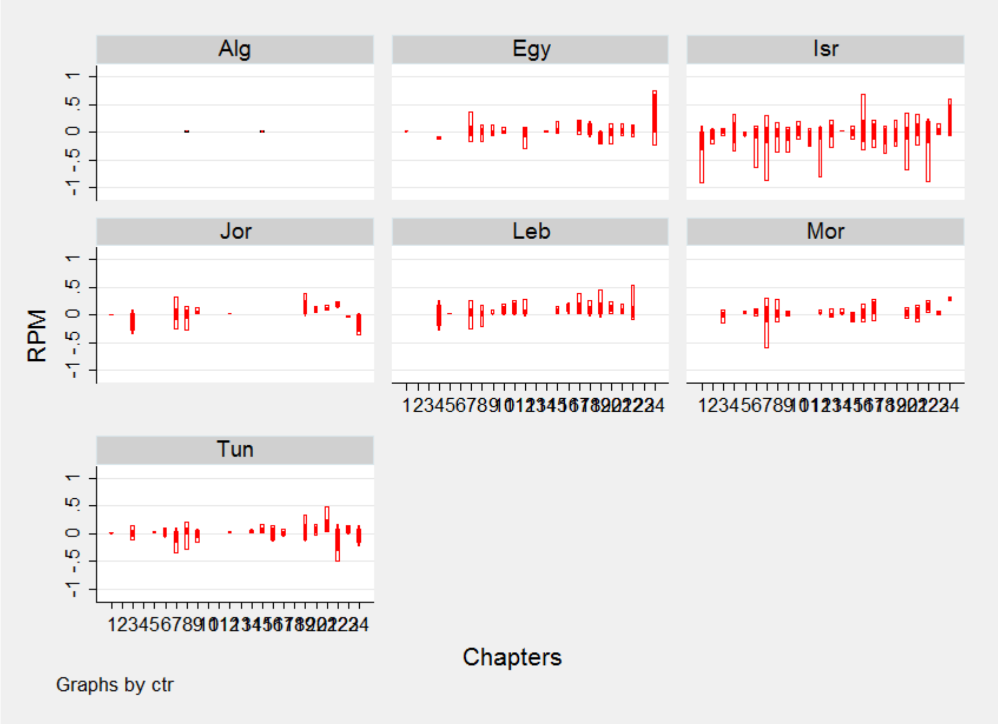


Source: Authors' own calculations on Comtrade

Figure A.3: **EU preferences towards SMCs (mean stdev min max)(before AA entry into force)**

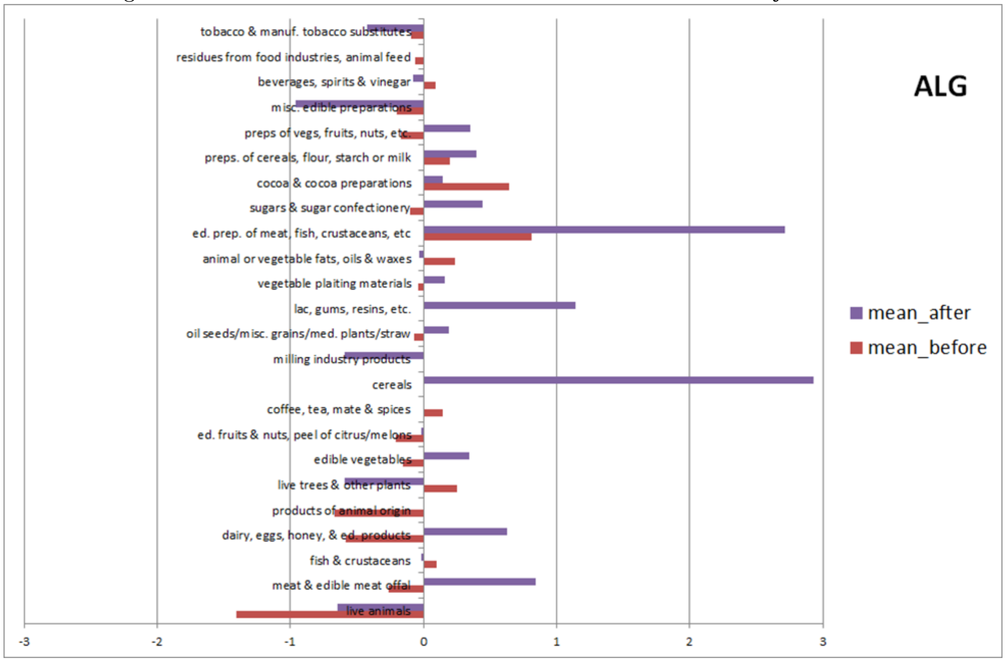
Source: Authors' own calculations on Comtrade

Figure A.4: **EU preferences towards SMCs (mean stdev min max)(after AA entry into force)**



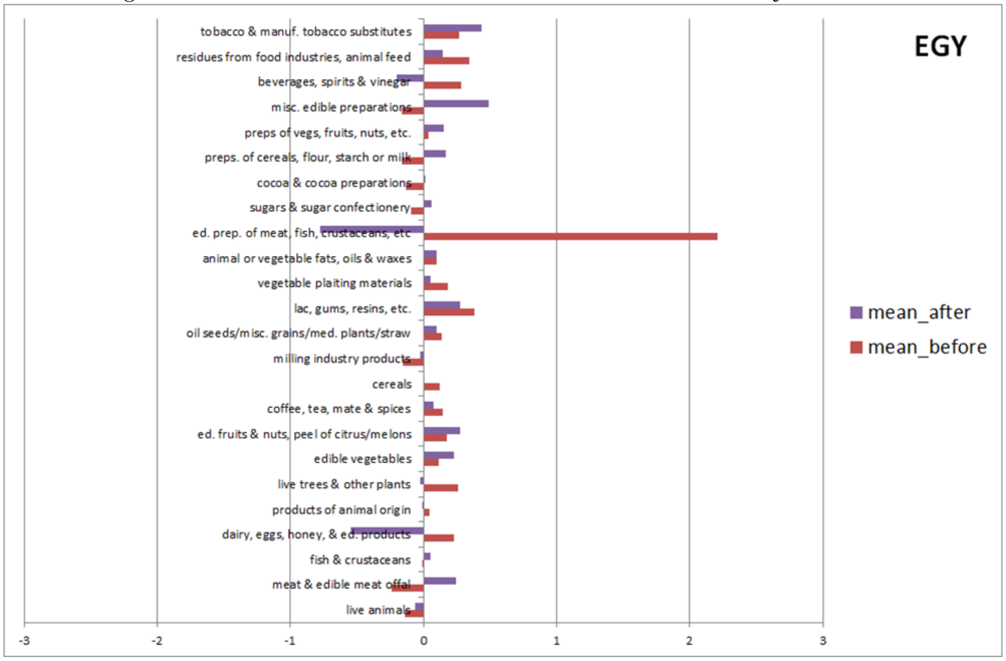
Source: Authors' own calculations on Comtrade

Figure A.5: **SMCs external trade: before and after AA Entry into force**



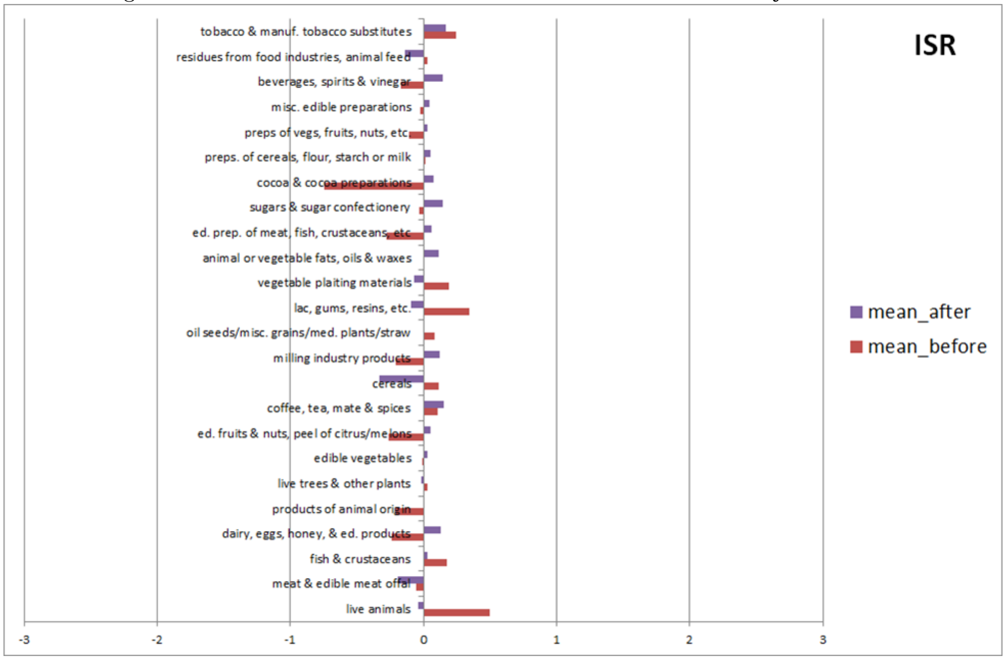
Source: Authors' own calculations on Comtrade

Figure A.6: **SMCs external trade: before and after AA Entry into force**



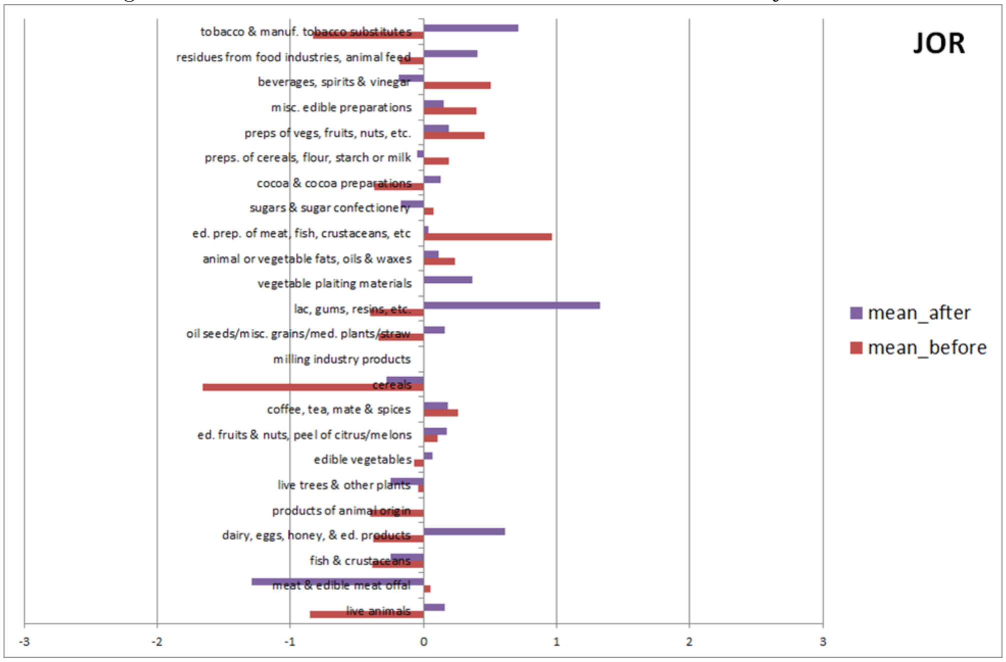
Source: Authors' own calculations on Comtrade

Figure A.7: **SMCs external trade: before and after AA Entry into force**



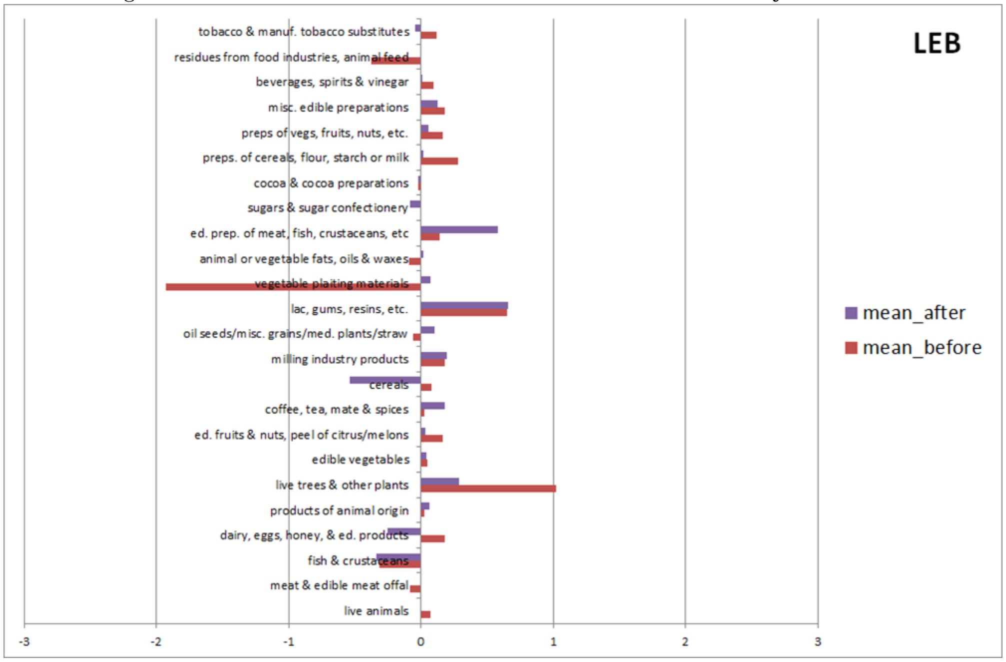
Source: Authors' own calculations on Comtrade

Figure A.8: **SMCs external trade: before and after AA Entry into force**



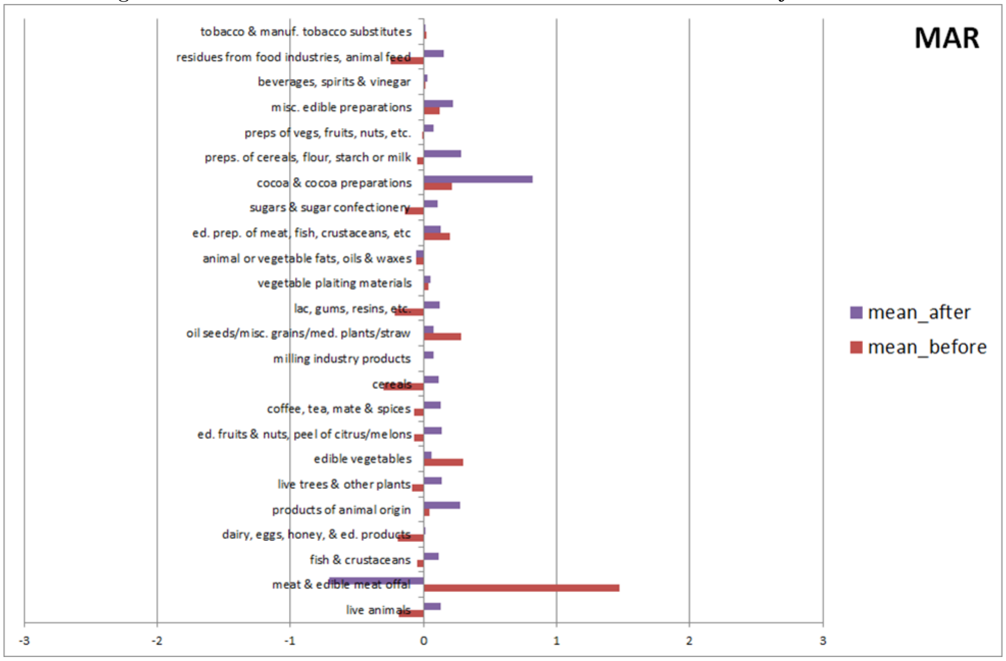
Source: Authors' own calculations on Comtrade

Figure A.9: **SMCs external trade: before and after AA Entry into force**



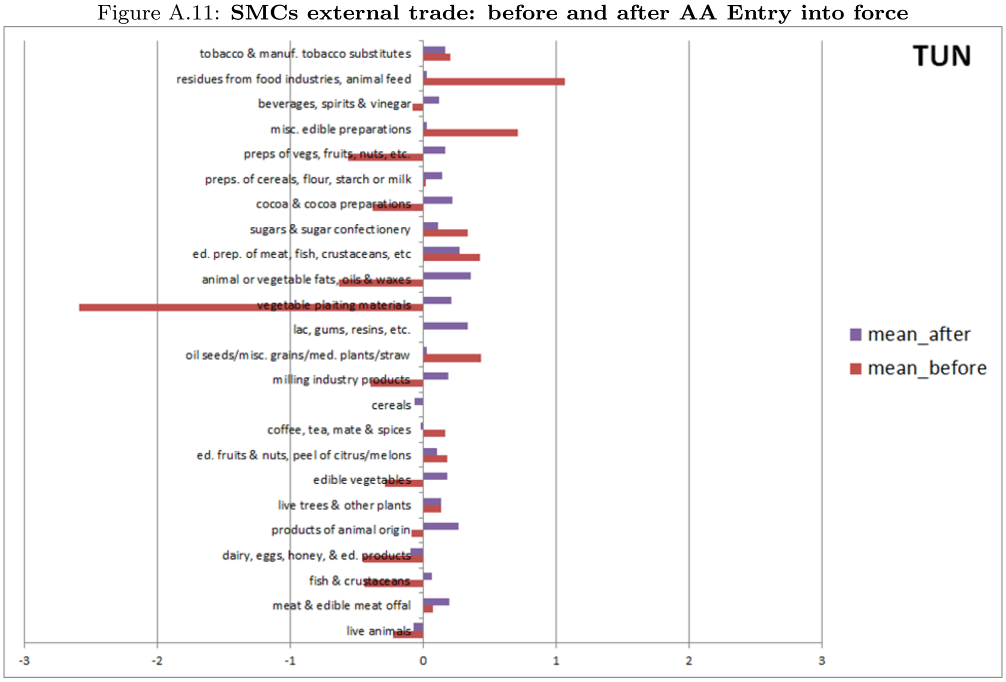
Source: Authors' own calculations on Comtrade

Figure A.10: **SMCs external trade: before and after AA Entry into force**



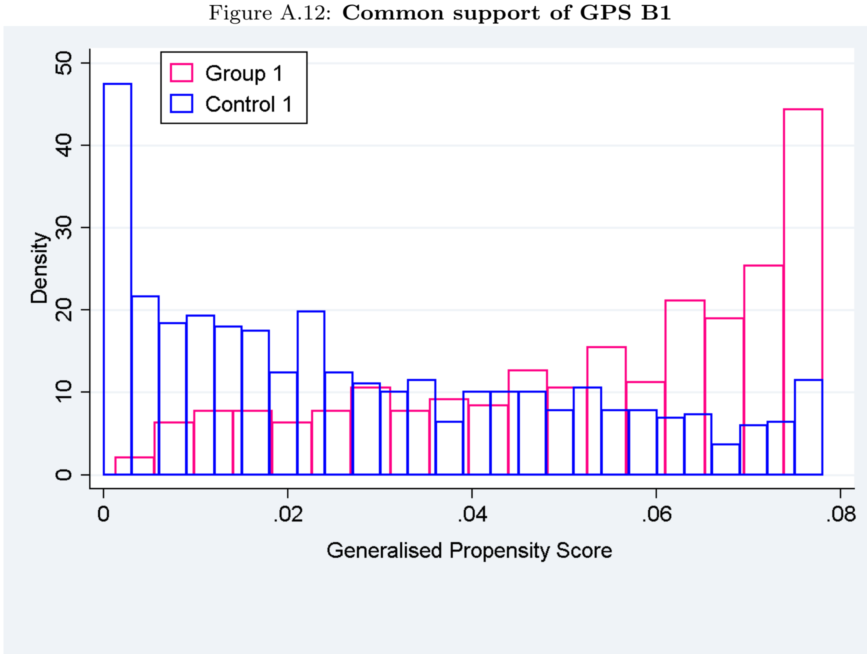
Source: Authors' own calculations on Comtrade

Figure A.11: **SMCs external trade: before and after AA Entry into force**



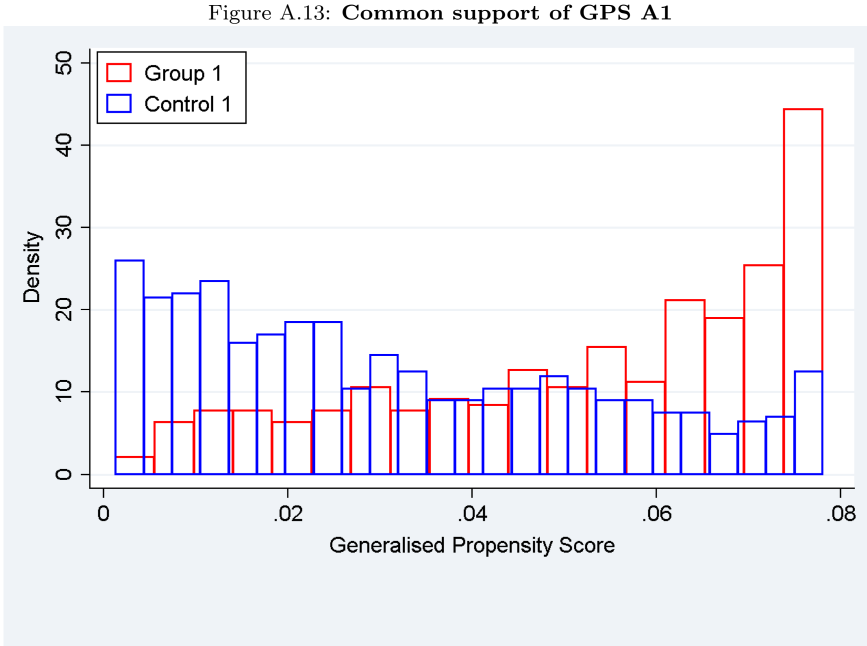
Source: Authors' own calculations on Comtrade

Figure A.12: **Common support of GPS B1**



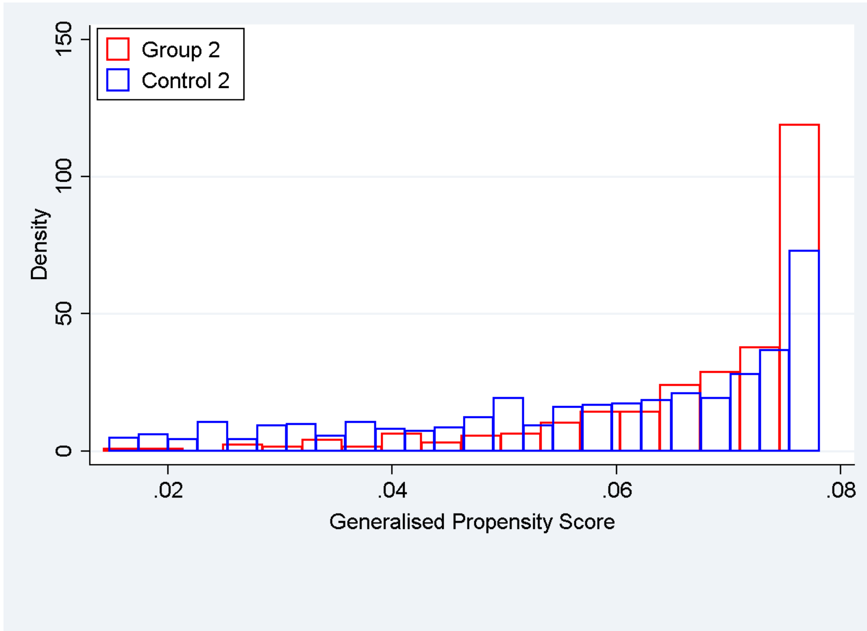
Source: Authors' calculations

Figure A.13: **Common support of GPS A1**



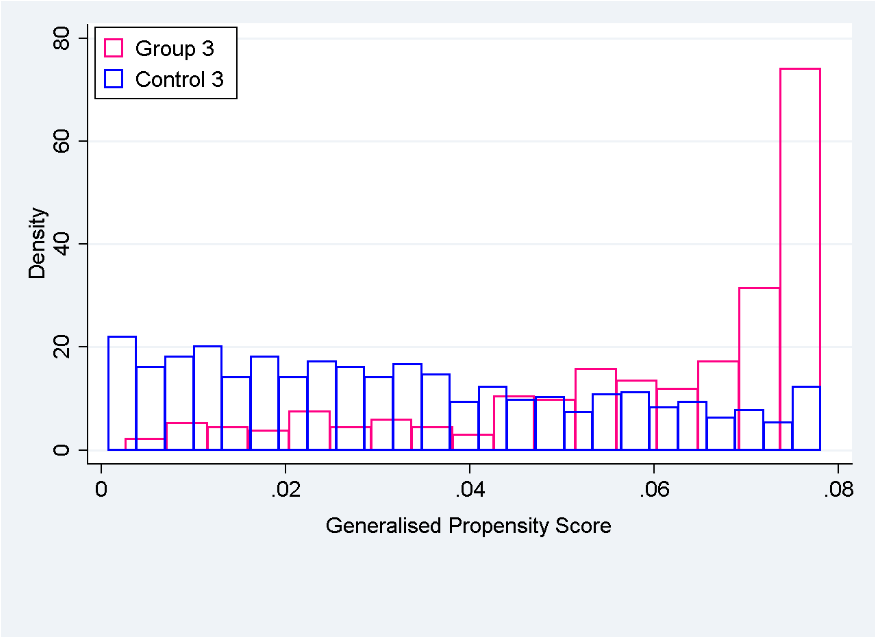
Source: Authors' calculations

Figure A.15: **Common support of GPS A2**



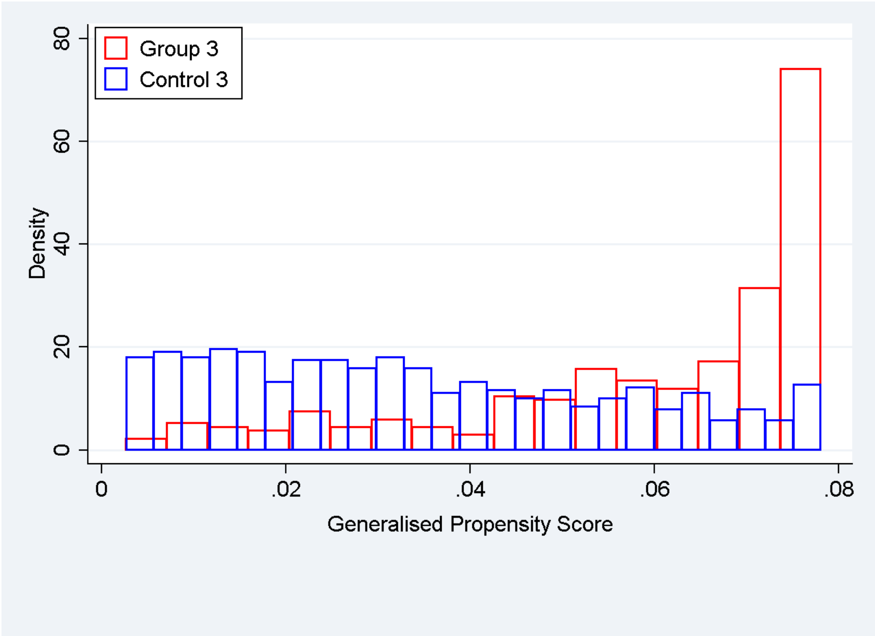
Source: Authors' calculations

Figure A.16: **Common support of GPS B3**



Source: Authors' calculations

Figure A.17: **Common support of GPS A3**



Source: Authors' calculations

FOODSECURE
FOR POLICIES THAT MATTER

The FOODSECURE project in a nutshell

Title	FOODSECURE – Exploring the future of global food and nutrition security
Funding scheme	7th framework program, theme Socioeconomic sciences and the humanities
Type of project	Large-scale collaborative research project
Project Coordinator	Hans van Meijl (LEI Wageningen UR)
Scientific Coordinator	Joachim von Braun (ZEF, Center for Development Research, University of Bonn)
Duration	2012 - 2017 (60 months)
Short description	In the future, excessively high food prices may frequently reoccur, with severe impact on the poor and vulnerable. Given the long lead time of the social and technological solutions for a more stable food system, a long-term policy framework on global food and nutrition security is urgently needed. The general objective of the FOODSECURE project is to design effective and sustainable strategies for assessing and addressing the challenges of food and nutrition security. FOODSECURE provides a set of analytical instruments to experiment, analyse, and coordinate the effects of short and long term policies related to achieving food security. FOODSECURE impact lies in the knowledge base to support EU policy makers and other stakeholders in the design of consistent, coherent, long-term policy strategies for improving food and nutrition security.
EU Contribution	€ 8 million
Research team	19 partners from 13 countries

FOODSECURE project office
LEI Wageningen UR (University & Research centre)
Alexanderveld 5
The Hague, Netherlands

T +31 (0) 70 3358370
F +31 (0) 70 3358196
E foodsecure@wur.nl
I www.foodscecure.eu

This project is funded by the European Union under the 7th Research Framework Programme (theme SSH) Grant agreement no. 290693