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Heterogeneous Effects of Tariff and Nontariff Policy Barriers in General Equilibrium

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Heterogeneous Effects of Tariff and Nontariff Policy Barriers in General Equilibrium

February 29, 2016

Abstract

Most applied work in international economics treats trade policy (a) as a linear component of trade costs and (b) as an exogenous variable. This paper proposes a structural modelling approach that allows for the estimation of (possibly) non-parametric effects of trade policy using a propensity score method to account for the endogeneity bias of trade policy. The findings point to important nonlinear effects of tariff and nontariff policy. Specifically, they suggest that interdependencies between nontariff policy barriers and tariffs are an important determinant of the partial impact of a given policy change. Overall, trade policy changes seem to be effective only for low- and medium-tariff trade flows. In general equilibrium, loosening the linearity assumption translates to an increased heterogeneity of predicted trade effects with the mean and median effect being up to three times as large in a flexible, nonparametric specification.

Keywords: Trade policy; Gravity models; Non-parametric and Semi-parametric Methods; Propensity Score Methods.

JEL classification: F14, F13.

1 Introduction

The quantitative role of trade impediments for the volume cross-border business transactions is key to understand the economic value of countries' openness. The inception of the General Agreement of Tariffs and Trade and the foundation of the World Trade Organization had been unprecedented milestones in the process of the coordinated liberalization of policy barriers to cross-border business activity. While this general view –mainly based on mean-effects– is uncontested, there is relatively little profound knowledge of what trade-policy liberalization is actually worth in quantitative terms once we allow the effect to be heterogeneous. There is many reasons to believe that a given change in trade-policy affects a country in a different way depending on

its status quo: Starting from prohibitive trade barriers or removing an already very low tariff-rate are likely to have very different effects on economic outcomes. Also, removing a tariff while a restrictive import quota is in place should be less effective than doing so when there are no nontariff barriers. While these questions are highly relevant for policy-makers, most empirical frameworks in applied international economic work are not suited to answer these questions:

1. Leading assumptions in the literature on the quantitative role of (policy or non-policy) trade barriers for economic outcome are that trade costs can be linearly decomposed into their components. In these models, any nonlinear effects of trade costs on trade flows are solely stemming from partial and general equilibrium effects.¹
2. Most work quantifying the effects of tariff barriers takes those as exogenous.²
3. Where endogeneity of trade policy is considered, it is mostly only crudely measured in terms of a binary membership versus non-membership in free trade agreements (treating all such agreements and, thereby, nonmembers and nonmembers, respectively, the same).³
4. Where trade policy is concerned, it is largely focused on tariff barriers, while mostly ignoring nontariff issues.

This customary approach used in the vast majority of empirical studies on trade barriers in international economics may lead to biased quantitative effects of trade policy for two reasons. First, ignored nonlinear direct effects may lead to a downward bias in the degree of heterogeneity of total effects of trade policy across countries and country pairs (as well as sectors). Second, ignored endogenous effects may lead to a bias of the magnitude of average as well as local (country-, country-pair-, or sector-specific) effects of policy. This paper puts forward a quantitative analysis which is aimed at addressing these two issues in a unified framework. In an effort to quantify marginal effects of potentially endogenous tariff and nontariff policy barriers to trade on economic outcome, we provide evidence of a large degree of nonlinearity of the partial effects of nontariff barriers and, to a lesser extent, also of tariff barriers. We allude to the reasons for such nonlinearities in the subsequent section. This usually ignored nonlinearity

¹See the seminal work by Anderson and van Wincoop (2002) and the subsequent work inspired by their methodology. Clearly, there are exceptions from this practice. E.g., Eaton and Kortum (2002) and Henderson and Millimet (2008) implicitly and explicitly, respectively, consider nonparametric effects of geography, both finding evidence of log-linearity about partial effects.

²See Ossa (2015), Caliendo and Parro (2014).

³See Baier and Bergstrand, 2007, 2009, though, with the exception of Egger, Larch, Staub, and Winkelmann, 2012, mostly in reduced-form settings

and heterogeneity adds substantially to the heterogeneity of total effects of marginal changes in trade policy.

The identification of heterogeneous partial responses of trade flows to endogenous tariff and nontariff trade-policy barriers in this paper rests on two pillars. One is the notion of weak unconfoundedness of trade-policy effects on trade flows conditional on all fundamental drivers of trade flows (productivity, factor endowments, number of households, etc.), as suggested by general equilibrium work on trade policy (see, e.g., Bond, 1988; Bond and Syropoulos, 1996; Bagwell and Staiger, 1997, 2004; Bond, Riezman, and Syropoulos, 2004). The second one is the flexible – parametric or nonparametric – relationship between trade-policy barriers and trade flows. These two ingredients support the use of a generalized-propensity-score approach to estimating the flexible behavioral response of interest. The approach is fully integrated into a customary general equilibrium model of bilateral trade with multiple sectors. There is an invaluable benefit from this integration along two lines. First, such a model helps decomposing data on trade flows in a way so that – prior to estimation and in conjunction with available data – all exogenous drivers of bilateral trade flows can be isolated. As a consequence, tariff and nontariff trade-policy variables may be conditioned on (filtered through) those variables, and residual terms which are unrelated to fundamental drivers of trade may be isolated. The latter may be used to assess the exogenous relationship between trade-policy variables and trade flows in a flexible, second step.

The main findings of the present study may be summarized as follows. We find substantial evidence for heterogeneous effects of policy barriers. A reduction in tariffs is only effective for the range of low and medium tariffs as well as for rather low values of nontariff barriers. In general, neglecting the interdependencies between tariff and nontariff policy barriers severely biases estimated changes in trade flows. High nontariff barriers are especially relevant in determining trade costs if at the same time medium and low levels of tariffs are in place.

Accounting for these heterogeneous effects in general equilibrium translates to an increased range (hence, additional heterogeneity) of predicted trade flow changes and rises the mean effect by a factor of around three as compared to traditional estimates based on a linear instead of a nonparametric specification. Finally, the general equilibrium exercise reveals that the restriction to linearity is especially misleading for high-tariff trade flows and high nontariff-barrier trade flows.

2 Theoretical background

Standard trade theory rarely models bilateral trade costs explicitly but treat them as an exogenous parameter. This paper is geared towards scrutinizing this assumption and allowing trade costs to be endogenous and of arbitrary functional form. In order to be able to evaluate and isolate the effect of this specific experiment, it is important to stay as close as possible to standard trade theory apart from the treatment of trade costs. For this reason we present the theoretical framework in two steps. First, we introduce a very stylized, general model of trade with its ingredients taken from the shelf of available trade theories. Having established that framework, we take a deeper look into the bilateral trade costs component.

2.1 A stylized trade model

To set up a theoretical frame for the analysis conducted in this paper, we aim at constructing a theoretical trade model that is as general as possible to nest most prevalent theoretical trade models but restrict ourselves by certain simplifying assumptions emphasizing that the contribution of the paper is an empirical one. Guided by the empirical reality, we employ a multi-country $i = 1, \dots, J$, multi-sector $s = 1, \dots, S$ setting.

On the supply side, to make our lives easier, we consider a world with one factor, labor L_i^s , that is immobile across sectors s and countries j and paid a wage rate W_i^s . Since this excludes any reallocation of factors of production, any equilibrium changes should be interpreted as short-run equilibrium effects. The value of production in each sector is assumed to be given by⁴

$$Y_i^s = W_i^s L_i^s. \quad (1)$$

On the demand side, we assume that a constant share β_j^s of income is spent on each sector. All income is derived from labor income, $E_j = \sum_{s=1}^S W_j^s L_j^s$. Within each sector, the allocation of expenditure towards country i is determined by

$$\pi_{ij}^s = \frac{A_i^s C_{ij}^s}{\sum_{k=1}^J A_k^s C_{kj}^s}, \quad (2)$$

where A_i^s denote any exporter-sector-specific components and C_{ij}^s denote country-pair-sector-specific components. Note that this specification guarantees the accounting identity $\sum_{i=1}^J \pi_{ij}^s = 1$ to hold.

In the remainder, we will broadly associate C_{ij}^s with a trade-cost component, noting, however, that C_{ij}^s includes exogenous effects of (Armington-type) consumer preferences in country j towards goods s from i . Allowing for a preferably wide range of trade theories to be nested,

⁴Note that this specification excludes e.g. fixed costs being paid in the importing country.

the only requirement on the exporter-sector-specific component is a multiplicatively separable form of endogenous and exogenous determinants of exporter potential

$$A_i^s = F_i^s (W_i^s)^{\alpha_s}, \quad (3)$$

where W_i^s is the (endogenous) country-sector-specific wage rate and F_i^s is the (exogenous) exporter-sector-specific fundamental driver of trade. These fundamentals can comprise supply-side factors such as factor endowments, productivity parameters and measures of comparative advantage as well as demand-side factors such as preference parameters. Accordingly, the structural interpretation of the sector-specific elasticity α_s depends on the respective fundamental (and hence structural trade model) we have in mind. Finally, assuming balanced trade allows us to solve for the endogenous country-sector-specific wage rate using

$$W_i^s L_i^s = \sum_{j=1}^J \frac{F_i^s (W_i^s)^{\alpha_s} C_{ij}^s}{\sum_k F_k^s (W_k^s)^{\alpha_s} C_{ij}^s} \beta_j^s \sum_{s=1}^S L_j^s W_j^s. \quad (4)$$

The theoretical framework outlined above nests (structurally) very different trade models. Among them, versions of Melitz's (2003) model of heterogeneous firms, Eaton and Kortum's (2002) model of Ricardian comparative advantage and Armington's (1969) model of national product differentiation - arguably, the most successful models in international economics.

2.2 A different view on trade costs

We leave standard theory in our treatment of bilateral trade costs. For simplicity, let us assume that the trade elasticity corresponds to α_s ⁵, leaving us with

$$c_{ij}^s = \alpha_s d_{ij}^s(m_{ij}^s, u_{ij}^s), \quad (5)$$

where we introduce the convention of denoting the log of any variable A by its lower-case counterpart a . We will refer to $d_{ij}^s(\cdot)$ as trade costs and assume it to be an arbitrary function of an (endogenous) policy vector $m_{ij}^s = (\tau_{ij}^s, \eta_j^s)$ and a vector of (exogenous) residual trade costs u_{ij}^s . Staying within the framework of standard trade theory introduced before, eventually, trade policy must be a function of the exogenous parameters of the model:⁶

$$m_{ij}^s = g_s(\mathbf{f}, \mathbf{u}, \boldsymbol{\alpha}), \quad (6)$$

⁵This simplification is especially useful to reduce the data requirements for the empirical analysis. However, our model does no longer generalize to a version of Melitz's (2003) model under this additional assumption. A modification that relaxes that assumption can be found in the appendix.

⁶Clearly, this specification excludes any influence of the demand parameter β_j^s on the determination of trade policy.

where $\mathbf{f} = (f_i^s) \forall i, s$, $\mathbf{u} = (u_{ij}^s) \forall i, j, s$ and $\boldsymbol{\alpha} = (\alpha_s) \forall s$. This specification allows the determination of trade policy to be a complex function of a multiplicity of factors: Fundamentals of each economy such as factor endowments determine trade policy as well as exogenous bilateral parameters such as a common border of trading partners. By allowing the policy determination to be sector specific we account for differences in policy determination across sectors. Finally, the trade elasticity itself can influence trade policy. Note that, by the bivariate structure of trade policy and its joint determination, we allow for any complementarity or substitutability between these variables. Assuming additive separability between exogenous and endogenous variables, the model of interest is given by

$$d_{ij}^s = h(m_{ij}^s) + \gamma' u_{ij}^s. \quad (7)$$

Clearly, tariff barriers do have a linear impact on trade costs via their ad-valorem price effect. Still there are several reasons to allow for additional flexibility beyond this linear effect, that we want to illustrate in what follows.

2.2.1 Nonlinear partial effects of trade costs in the literature

Two fundamental theoretical reasons for nonlinear partial effects of trade costs in general and trade policy in specific can be distinguished:

- the trade-cost function is nonlinear in its arguments, or
- the price-elasticity of demand is not constant.

For a conditional econometric assessment of partial effects, the two are not distinguishable, and which reason is relevant only matters for the analysis of theory-consistent total effects of trade-cost changes.

Nonlinear partial effects of trade policy due to nonlinear trade-cost functions

With a constant, homogeneous price elasticity of demand, partial responses to (identical) changes in trade costs are the same across countries and country pairs within a sector, but trade-policy and other variables may affect trade costs in a log-nonlinear way. For nontariff barriers a log-nonlinear relationship to trade costs seems ex ante relatively more natural than for tariff barriers. The reason is that the measurement of nontariff barriers and their nature and role are less understood than those of tariff barriers. However, even tariff barriers and other ad-valorem costs may affect trade costs log-nonlinearly, according to the literature for two fundamental reasons, one pertaining to high and the other one to low tariff rates.

At very high tariff rates, the claim is – and evidence confirms it – that importers and exporters find it desirable to engage in standard (illegal) tax-avoidance strategies. In that regard, earlier work pointed to, e.g., the importance of the misdeclaration of products, at least in countries with a weak governance (see, e.g., Fisman and Wei, 2004; Javorcik and Narciso, 2008).

At very low tariff rates, earlier research points to a lower partial impact than on average, and the reason given is that low, positive applied tariffs often reflect unused preferential market access (see Grether and Olarreaga, 1998; Carpenter and Lendle, 2010; Fugazza and Nicita, 2013). The respective literature attributes the log-nonlinear relationship between (one-plus)tariff rates and trade flows to administrative and other nontariff aspects of tariff regimes. Grether and Olarreaga (1998) conclude that tariffs only matter if they are higher than three percentage points.

Another important channel for nonlinear partial effects of tariff policy is through second order effects. If there is a role for uncertainty in trade costs, this effect is likely to be dependent on the tariff and nontariff barrier level. Second order effects may similarly arise through the interdependences and substitutabilities between tariff and nontariff barriers: The level of an (endogenously determined) tariff may depend on the current level of nontariff barriers.

Work on the natural (geographical) part of trade costs suggests that non-policy barriers affect demand in a log-nonlinear way at short distance (see Hillberry and Hummels, 2008), while the relationship appears indistinguishable from being log-linear at larger distances (see Eaton and Kortum, 2002; Henderson and Millimet, 2008).

Nonlinear partial effects of trade policy due to variable price elasticities of demand

Another strand of work explores the role of variable price elasticities of demand (see, e.g., Feenstra and Kee, 2008; Feenstra and Weinstein, 2010; Mrázová and Neary, 2013; Novy, 2013; Mrázová, Neary, and Parenti 2015). While some of this literature is mostly concerned with the variable relationship of costs to markups and producer prices to demand volume, an immediate implication is that the partial effect of trade costs to trade flows is heterogeneous in samples and the population of data as well (see Novy, 2013, for an illustration with translog preferences), of demand elasticities are variable.

2.3 General equilibrium effects of changes in trade costs

Having a specific trade cost function at hand the effect of any trade policy intervention, lets say from $m = (m_{ij}^s, \dots) \forall i, j, s$ to m' translates to a change of trade costs from d to d' . Following Dekle, Eaton and Kortum (2007), we can obtain the general equilibrium effect of a change in

trade costs using:

$$\hat{W}_i^s = \frac{1}{W_i^s L_i^s} \sum_j \left[\pi_{ij}^s E_j^s \left[\frac{\sum_s \hat{W}_j^s W_j^s L_j^s}{\sum_s W_j^s L_j^s} \right] \frac{(\hat{W}_i^s \hat{D}_{ij}^s)^{\alpha_s}}{\sum_k \pi_{kj}^s (\hat{W}_k^s \hat{D}_{kj}^s)^{\alpha_s}} \right], \quad (8)$$

where for any generic variable $\hat{A} = A'/A$.

3 Empirics

3.1 Econometric methodology

Following the theoretical model of section 2, an empirical specification of bilateral trade flows can be decomposed in an exporter-sector-specific effect, an importer-sector-specific effect and the bilateral trade-costs component

$$x_{ij}^s = a_i^s + b_j^s + \alpha_s(h(m_{ij}^s) + \gamma' u_{ij}^s) + \epsilon_{ij}^s. \quad (9)$$

However, the simultaneous determination of trade flows and trade policy is creating an endogeneity bias, $E(\epsilon_{ij}^s | m_{ij}^s) \neq 0$, that does not allow for a standard semi-parametric estimation of (9).⁷ Instead, we build on an econometric strategy –the generalized propensity score– that models a reduced form determining the policy vector m_{ij}^s in a first step and, in a second step, uses this reduced form to disentangle the causal impact of trade policy on bilateral trade flows. This methodology based in the treatment effect literature was first introduced by Hirano and Imbens (2004) and Imai and van Dyk (2004). For each country-sector pair indexed by i, j, s , let us define a potential trade flow outcome $x_{ij}^s(m)$ where $m \in \mathcal{M}$ denotes a bivariate vector of potential treatments of trade policy. In contrast to the potential treatment vector m , we refer to the actually observed level of treatment as m_{ij}^s . To identify the effect of treatment on outcome, hence the causal effect of some trade policy on trade flows, we rely on two assumptions.⁸

Assumption 1 (Stable unit treatment value assumption) *Conditional on some vector of observed covariates q_{ij}^s , the distribution of potential outcomes for one unit is independent of the (potential) treatment status of any other unit.*

Assumption 2 (Weak unconfoundedness assumption) *Any potential trade flow outcome is independent of the actual treatment conditional on some vector of observed covariates q_{ij}^s :*

$$x_{ij}^s(m) \perp m_{ij}^s | q_{ij}^s \quad \forall m \in \mathcal{M}. \quad (10)$$

⁷See e.g. Robinson (1988).

⁸Compare Imai and van Dyk (2004), Rubin (1980, 1990) and Hirano and Imbens (2004).

For these assumptions to be met, it is core to find a suitable vector of covariates q_{ij}^s . From the theoretical trade model of section 2, we know that all variables that potentially determine both trade flows and trade policy are $(\mathbf{f}, \mathbf{u}, \boldsymbol{\alpha})$. To maintain empirical tractability, we restrict the vector of covariates q_{ij}^s to contain only the respective variables indexed by i, j, s , (f_i^s, f_j^s, u_{ij}^s) . By this restriction we capture all first-order determinants of trade policy while we summarize second order determinants by their respective means, $(\bar{f}_{-i,-j}^s, \bar{u}_{-i,-j}^s)$. To account for sectoral differences in the policy determination function as well as to account for sector-specific trade elasticities, we allow for a sector specific constant, (i^s) . In what follows, we argue that – guided by theoretical trade theory – both assumptions are met given a vector of covariates $q_{ij}^s = (f_i^s, f_j^s, u_{ij}^s, \bar{f}_{-i,-j}^s, \bar{u}_{-i,-j}^s, i^s)$. Let $r(m, q) = f(m|q)$ be the conditional density of policy treatment m and refer to its realization $R=r(M, Q)$ as the generalized propensity score (GPS). It can be shown that,⁹

$$\begin{aligned} E(x_{ij}^s(m)|R = r, M = m) &= E(x_{ij}^s(m)|r(m, q_{ij}^s) = r) = a_i^s + b_j^s + \alpha_s(k(m, r(m, q_{ij}^s)) + \gamma' u_{ij}^s). \\ E(x_{ij}^s(m)) &= a_i^s + b_j^s + \alpha_s(E(k(m, r(m, q_{ij}^s))) + \gamma' u_{ij}^s). \end{aligned} \quad (11)$$

Hence, using the generalized propensity score, we are able to remove any bias associated with the joint determination of trade policy and trade flows and obtain a consistent estimate of the average causal effect of a given trade policy vector m . Note that the first part of (11) is a reduced form having no causal interpretation.

3.2 Data

The main source of information is WITS database on bilateral trade flows at the 4-digit level for the year 2006 as well as TRAINS database on tariffs for the year 2006. Data on import demand elasticities and ad-valorem equivalents of nontariff barriers are sourced from Kee, Nicita and Olarreaga (2008) and Kee, Nicita and Olarreaga (2009), respectively.¹⁰ Data on exogenous trade barriers is taken from CEPII. Finally, UNIDO provides us with sector-level production data.

3.2.1 Trade policy variables

We consider a bivariate vector of trade policy consisting of a tariff policy measure and a nontariff barrier measure. Using the ad-valorem values of tariff ($AVE_Tariff_{ij}^s$) and nontariff barriers ($AVE_NTB_j^s$) we define $\tau_{ij}^s = \log(1 + AVE_Tariff_{ij}^s)$ and $\eta_j^s = \log(1 + AVE_NTB_j^s)$.

⁹Refer to theorem 3.1 in Hirano and Imbens (2004) for a proof.

¹⁰While these data come at the 6-digit level, we aggregate them up on 4-digit level consistently with their estimation methodology. See their online appendix for more details.

3.2.2 Trade fundamentals

In order to obtain estimates for the fundamentals of the model, we run a fixed effect regression of bilateral exports with exporter-sector and importer-sector fixed effects only

$$x_{ij}^s = a_i^s + o_j^s + z_{ij}^s, \quad (12)$$

where $o_j^s = b_j^s + h(\eta_j^s)$ and $z_{ij}^s = z(c_{ij}^s, \eta_j^s)$. In our data, nontariff barriers η_j^s are bilateral only to the extent that they are assumed to be unity for a country's sales to itself but they are importer-specific for all other bilateral trade flows. Hence, the flexible trade cost function contains importer specific terms that are captured in o_j^s . Using $B_j^s = \frac{E_j^s}{\sum_{k=1}^J A_k^s C_{kj}^s}$ together with $C_{ij}^s = \exp(z_{ij}^s + h(\eta_j^s))$ for all $i \neq j$ and $C_{ij}^s = 1$ for $i = j$, we can isolate

$$B_j^s = \frac{O_j^s}{A_j^s} \left[\frac{E_j^s}{O_j^s} - \sum_{k \neq j} A_k^s Z_{kj}^s \right], \quad (13)$$

where we proxy sector level expenditure by $E_j^s = (Y_j^s - \sum_{k \neq j} X_{jk}^s) + \sum_{k \neq j} X_{kj}^s$ using data on bilateral trade flows X_{ij}^s and sector-level production Y_j^s . With an estimate of A_i^s and B_j^s at hand we can back out the residual $C_{ij}^s = X_{ij}^s / A_i^s B_j^s$. Bilateral trade costs $D_{ij}^s = (C_{ij}^s)^{1/\alpha_s}$ as well as $F_i^s = A_i^s / (W_i^s)^{\alpha_s}$ are readily calculated using data on factor prices W_i^s and importer price elasticities α_s . We use fob unit values to proxy for factor prices W_i^s .¹¹ Note that the aim of disentangling the exporter fixed effect is to isolate the endogenous part W_i^s from the exogenous part F_i^s . Thinking of fob unit values as producer prices (consumer prices abstracting from trade costs), they are proportional to factor prices in most models and allow to obtain an estimate of fundamentals that is completely exogenous.¹² Figure 1 shows the distribution of estimated exporters-sector specific fixed effects and the distribution of the fundamentals calculated thereof.

3.2.3 Exogenous trade barriers

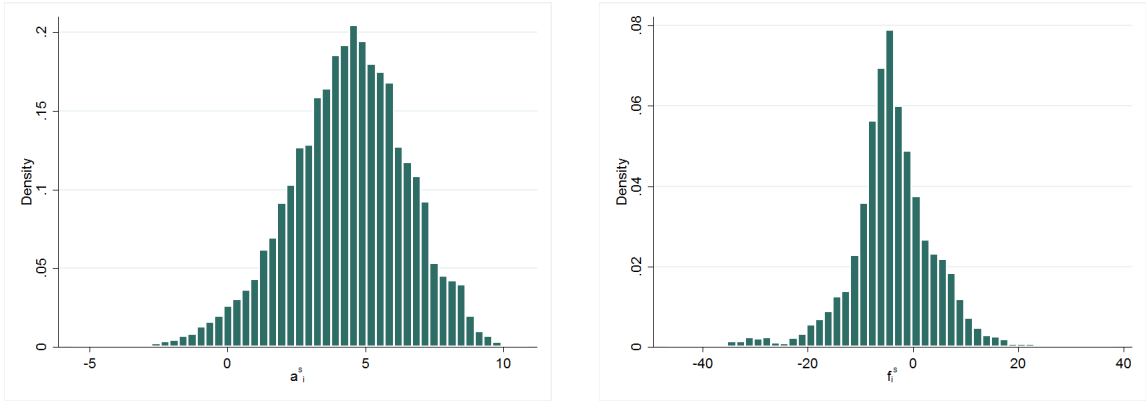
To account for exogenous trade barriers u_{ij}^s we use data on distance together with dummies for common borders, common languages (official and ethnological), common colonial past and a dummy for countries that used to be one country. Finally, we account for being a member of the same customs union since these matter in the determination of trade policy.¹³

¹¹We obtain fob unit prices by dividing a country's trade value in a given sector by trade quantity.

¹²A table relating producer prices to factor prices for all models can be found in the appendix.

¹³Customs unions considered are Comunidad Andina, CARICOM, East-African Customs Union, European Union, Mercosur, South-African Customs Union, Lichtenstein-Switzerland.

Figure 1: Results for fixed-effects regression.



(a) Histogram of estimated exporter-sector specific effects.

(b) Histogram of exporter-sector specific fundamentals calculated from producer prices and exporter-specific fixed effects.

3.2.4 Summary statistics

We use the full set of available information to estimate the fixed effects regression (12) which is comprised of the universe of bilateral trade data in 2006.¹⁴ Unfortunately, only a subset of these data can be used for estimating (11). On one hand, we are constrained by data availability, mainly by the incompleteness of sectoral production data and the availability of elasticity as well as nontariff trade barrier measures. Also, in order to match both f_i^s and f_j^s to a given bilateral trade flow x_{ij}^s , we need the importer to be an exporter of the same good. On the other hand, the econometric methodology employed excludes any mass points in the treatment variables.¹⁵ Given the wide range of zeros in both tariff and nontariff policy barriers, we drop these observations. Hence, any results are to be treated as valid for bilateral trade flows with positive policy barriers only. Summary statistics for both samples can be found in table 1. Clearly, average tariff and nontariff barriers are higher in the restricted sample where all zero barriers have been dropped. Also, the average trade volume observed is higher in the restricted sample because of better GDP data coverage in larger countries and sectors. In the restricted sample we observe 145 exporting countries, 48 importers and 115 manufacturing sectors.

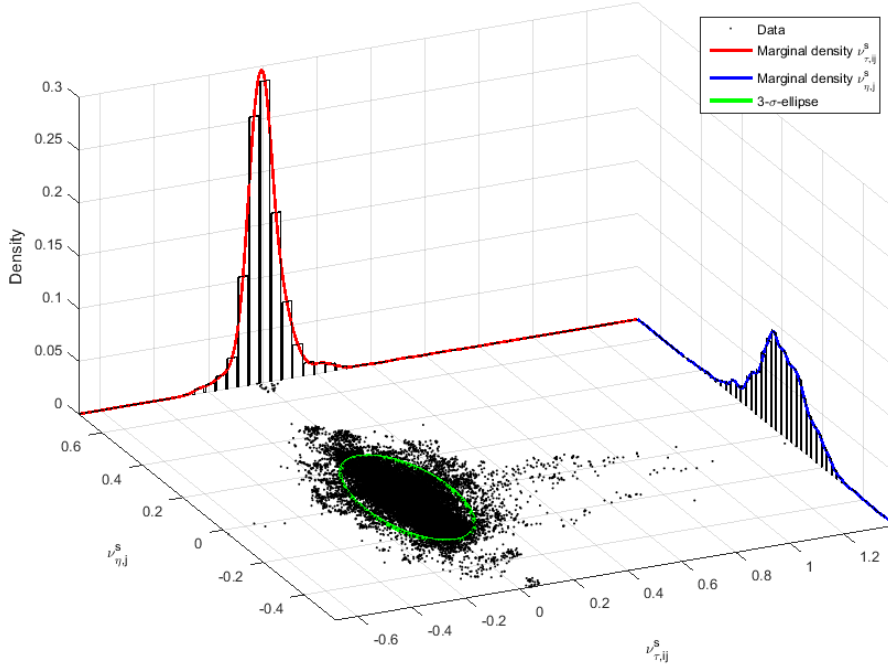
Table 1: Summary statistics.

Variable	Mean	Std. Dev.	N	Mean	Std. Dev.	N
log(Exports)	4.53	3.589	808127	6.483	3.19	33539
log(Quantity)	9.18	4.356	784905	11.562	3.757	33423
Elasticity	-2.428	2.129	455120	-2.41	1.523	33539
log(GDP)	20.661	2.358	453071	14.548	1.863	22018
Contiguity	0.061	0.239	745844	0.043	0.204	33538
Common lang. (off)	0.188	0.391	745844	0.131	0.338	33538
Common lang. (ethn.)	0.202	0.401	745844	0.165	0.371	33538
Colony	0.047	0.212	745844	0.051	0.22	33538
Common colony	0.069	0.254	745844	0.034	0.182	33538
Current colony	0.004	0.064	745844	0	0.009	33538
Colony (after 1945)	0.029	0.167	745844	0.027	0.161	33538
Same Country	0.026	0.16	745844	0.009	0.092	33538
log(Distance)	8.310	1.016	745844	8.683	0.801	33538
τ_{ij}^s	0.058	0.09	494402	0.104	0.105	33539
η_j^s	0.101	0.166	455120	0.229	0.18	33539
	Full sample			Restricted sample		

Table 2: Report of Akaike information criteria, adjusted R^2 and the list of covariates for the first stage regression.

Polynomial	τ_{ij}^s		η_j^s	
	adj. R^2	AIC	adj. R^2	AIC
2	31.68%	-68535.62	48.89%	-42192.72
3	31.75%	-68396.16	49.05%	-42312.04
4	32.05%	-68532.03	49.12%	-42269.87
5	33.15%	-69078.73	49.22%	-42334.75
6	33.38%	-69195.58	49.22%	-42331.03
7	33.39%	-69191.09	49.23%	-42342.81
8	33.68%	-69337.93	49.56%	-42562.49
9	33.74%	-69383.00	49.69%	-42639.72
10	33.79%	-69409.55	49.73%	-42679.38
11	33.86%	-69467.83	49.77%	-42699.70
12	33.93%	-69476.93	49.78%	-42710.01
13	33.93%	-69491.37	49.78%	-42710.01
14	33.93%	-69537.42	49.83%	-42803.80
15	33.94%	-69497.45	49.83%	-42803.80
16	33.94%	-69491.97	49.83%	-42803.80
17	33.97%	-69609.85	49.87%	-42761.55
18	33.97%	-69495.94	49.87%	-42761.55
Covariates	$\hat{f}_i^s, \hat{f}_j^s, \log(\text{distance}), \text{contiguity},$ common ethno. lang., colony, common colonizer, common off. lang., current colony, colony(1945), same country, $\bar{f}_{-i,-j}^s, (\text{mean})\log(\text{distance}),$		$\hat{f}_j^s, \log(\text{distance}), \text{contiguity},$ common ethno. lang., colony, common colonizer, common off. lang., current colony, colony(1945), same country, $\bar{f}_{-i,-j}^s, (\text{mean})\log(\text{distance}),$	
Interaction terms	Yes		Yes	
Sector & CU FE ¹⁶	Yes		Yes	
Observations	33538		33538	

Figure 2: Bivariate distribution of ν_{ij}^s .



3.3 Estimation

In a first step, we obtain an estimate for the conditional bivariate treatment $m_{ij}^s | q_{ij}^s$ by estimating a polynomial regression of the policy variables in m_{ij}^s on the covariates q_{ij}^s

$$\tau_{ij}^s = g_\tau(\hat{f}_i^s, \hat{f}_j^s, u_{ij}^s, \bar{f}_{-i,-j}^s, \bar{u}_{-i,-j}^s, i^s) + \nu_{\tau,ij}^s \quad (14)$$

$$\eta_j^s = g_\eta(\hat{f}_j^s, u_{ij}^s, \bar{f}_{-i,-j}^s, \bar{u}_{-i,-j}^s, i^s) + \nu_{\eta,j}^s, \quad (15)$$

with the degree of the polynomial being determined by an Akaike information criterion.¹⁷ Table 2 provides the information criteria and adjusted R^2 for the two regressions. Based on the information criterion we chose a polynomial of degree 17 and 14, respectively. Both regressions show a very high explanatory power, reassuring the choice of covariates. The residuals $\nu_{ij}^s = (\nu_{\tau,ij}^s, \nu_{\eta,j}^s)$ serve as an estimate for the conditional bivariate treatment whose distribution has to be estimated in a next step. Figure 2 plots the bivariate distribution of the estimated residuals. We estimate the conditional density of policy treatment by three different means. First, we assume a bivariate normal distribution and estimate the parameters of the distribution by maximum likelihood. While figure 2 shows that the distribution is (close) to being bell-shaped it is still a

¹⁴We use the year 2006 because this provides us with best coverage.

¹⁵See Imbens (1994) for a discussion.

¹⁷Note that we can not include any i-specific regressors in the second regression on j-specific η_j^s .

very restrictive assumption to assume the distributional family to be normal ex-ante. Therefore, we employ a more approach that estimates the distribution only under the shape constraint of being log-concave – with the multivariate normal distribution nested by the class of log-concave distributions.¹⁸ Since the multivariate normal distribution is nested by the class of log-concave distributions, we allow for more flexibility in the estimation. Finally, we allow for a maximum degree of flexibility by estimating the unconditional bivariate density of ν_{ij}^s following Li and Racine (2007).¹⁹ The estimated density serves as an estimate for the propensity function: for any potential treatment we might be interested in it allows to estimate a GPS, a scalar indicating the propensity for that specific treatment, for any country-sector pair in our data. Having obtained an estimate of this propensity function, it is indicated to check the common support property and drop any observations with extreme GPS values.²⁰ since extreme values of the GPS do not allow for a comparison across observations. We follow Flores, Flores-Lagunes, Gonzalez and Neumann (2012) in defining the common support and extend their methodology to multivariate treatments. Let us define 4×4 groups corresponding to all possible combinations of quartiles of the two policy treatment variables. We take the median value of the two policy treatments within each group and calculate the GPS corresponding to these 4 different potential treatment vectors for each observation. Any observations with a GPS corresponding to treatment vector k lying outside the range spanned by observations in group k is dropped. In total we drop 875 observations leaving us with a sample of 32'664 trade flows for the estimation.²¹

In order to assess the validity of the estimated GPS, we make use of the balancing property of the propensity score.²² By definition of the propensity score, for any observation with the same GPS, $r(m, q_{ij}^s)$ the probability that $M = m$ is independent of the covariates q_{ij}^s . Similarly to before, we create 3×3 groups corresponding to the combinations of tertiles of the two policy variables. For every covariate (including the fixed effects) we check the difference of the mean

¹⁸A function f on \mathcal{X}^d is log-concave if it is of the form $f(x) = \exp \phi(x)$, where $\phi(\cdot)$ is some concave function. The class of log-concave densities contains most of the commonly used parametric distributions, e.g., normal distribution, exponential distribution, logistic distribution, extreme value distribution, Laplace distribution, chi distribution, gamma distribution if the shape parameter is ≥ 1 , chi-squared distribution with a number of degrees of freedom of ≥ 2 , beta distribution if both shape parameters are ≥ 1 , Weibull distribution if the shape parameter is ≥ 1 . We heavily build on the work by Cule, Samworth, and Stewart (2008) who show how the MLE of a multivariate log-concave density can be computed.

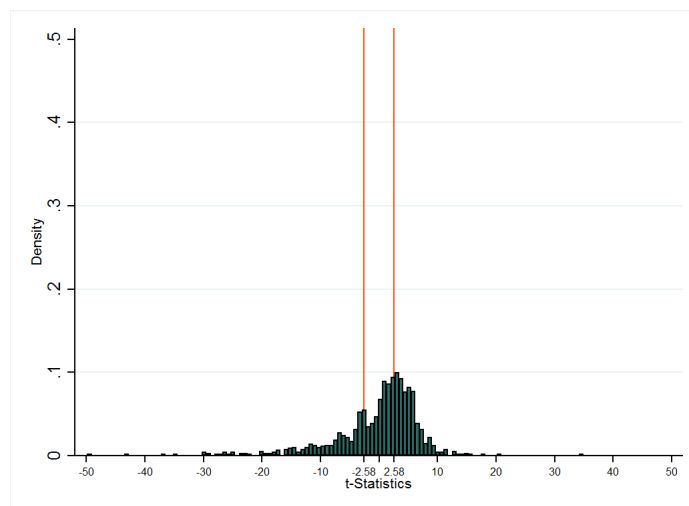
¹⁹We choose the bandwidth by likelihood cross-validation (compare Li and Racine (2007)).

²⁰Compare Dehejia and Wahba (2002), Imbens (2004) and Flores, Flores-Lagunes, Gonzalez and Neumann (2012)

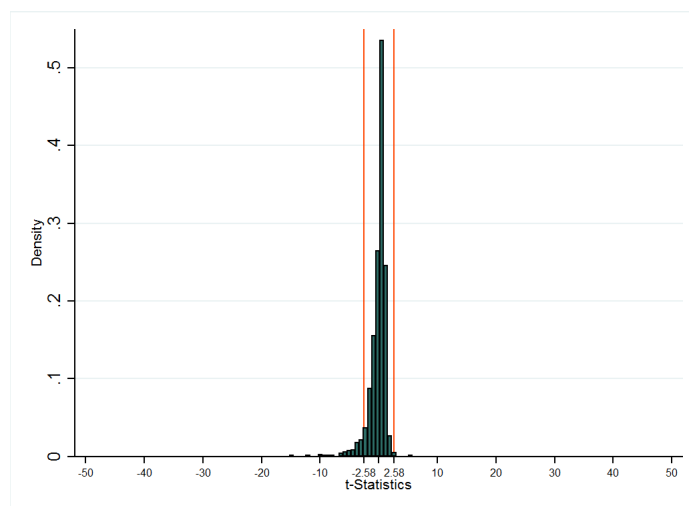
²¹Note that the common support restriction was conducted with the propensity function being estimated by the normality assumption in a first step. This is done to ensure the same sample being used for every specification.

²²Compare Hirano and Imbens (2004).

Figure 3: Distribution of t-statistics of equality-of-means test for all covariates.



(a) Without accounting for the GPS.



(b) Accounting for the GPS.

Table 3: t-statistics of equality-of-means test for non-fixed-effects covariates without accounting for the GPS.

Covariates	Group								
	1	2	3	4	5	6	7	8	9
f_i^s	-5.52	-14.09	-7.07	-11.40	-2.97	11.48	1.14	14.77	14.07
f_j^s	-17.53	-18.83	-9.21	-6.59	0.37	7.31	9.75	20.31	15.30
log(Distance)	-7.70	2.61	6.41	-9.50	-3.88	-0.77	1.47	3.39	7.41
Contiguity	0.61	-4.94	-4.07	-3.40	-0.63	4.45	0.58	2.01	5.55
Common lang. (off.)	3.56	1.88	-6.05	2.66	-1.33	-0.66	-0.78	0.57	0.18
Common lang. (ethno.)	3.53	-1.45	-6.62	4.15	-3.21	1.62	-2.73	0.85	4.03
Colony	2.09	2.11	-2.01	3.85	1.04	-4.58	5.79	0.43	-8.75
Common colonizer	9.12	2.11	-0.13	4.25	0.47	5.24	-12.58	-7.81	-0.52
Colony (1945)	4.32	3.75	-1.79	3.95	0.33	-0.92	1.68	-3.74	-7.57
Same country	-0.75	-2.27	-2.82	-0.21	-0.02	3.92	-1.96	0.64	3.72
$\bar{f}_{-i,-j}$	-6.47	-15.32	-10.06	-10.66	-1.99	11.73	2.56	17.51	13.30
mean(Distance)	7.70	-2.60	-6.41	9.51	3.88	0.77	-1.48	-3.40	-7.42
CAN	7.52	13.20	11.64	-29.43	-14.06	-3.55	-1.31	2.15	11.17
CARICOM	3.49	3.16	3.48	3.24	1.17	3.24	-9.87	-11.27	3.41
EAC	5.08	5.07	1.92	4.72	4.48	-2.19	4.96	1.95	-26.19
EU	-11.22	9.65	4.06	15.38	12.56	-43.00	34.55	11.00	-34.61
MERCOSUR	9.13	7.95	8.08	-1.83	-0.26	8.34	-12.69	-13.64	-5.34
SACU	2.66	-2.47	2.66	2.47	2.47	2.47	-15.25	2.59	2.60
LICH	-7.25	0.96	0.96	0.90	0.90	0.90	0.94	0.94	0.94

Table 4: t-statistics of equality-of-means test for non-fixed-effects covariates accounting for the GPS (unconditional density).

Covariates	Group								
	1	2	3	4	5	6	7	8	9
f_i^s	-0.47	-1.24	-2.19	-1.57	0.24	1.82	-0.63	2.06	1.68
f_j^s	-2.11	-1.81	-2.41	-0.76	0.88	1.38	0.72	2.94	1.83
log(Distance)	-0.32	0.81	1.39	-1.06	-0.60	-0.24	-0.20	0.60	1.43
Contiguity	-0.20	-0.80	-0.69	-0.77	-0.33	0.63	0.11	0.28	0.61
Common lang. (off.)	0.50	0.37	-1.24	0.17	-0.40	-0.50	0.22	0.05	0.04
Common lang. (ethno.)	0.45	-0.26	-1.32	0.25	-0.61	-0.16	-0.15	0.14	0.74
Colony	0.39	0.37	-0.57	0.67	0.24	-0.65	0.91	-0.15	-1.59
Common colonizer	0.95	0.12	0.10	0.47	-0.05	0.61	-2.44	-1.38	-0.28
Colony (1945)	0.66	0.51	-0.25	0.63	0.15	-0.22	0.27	-0.87	-1.42
Same country	-0.20	-0.28	-0.36	-0.02	-0.01	0.54	-0.33	0.18	0.56
$\bar{f}_{-i,-j}$	-0.62	-1.41	-2.52	-1.45	0.46	1.91	-0.45	2.38	1.48
mean(Distance)	0.32	-0.81	-1.39	1.06	0.60	0.24	0.20	-0.60	-1.43
CAN	1.46	2.18	1.22	-4.36	-2.23	-0.91	-1.18	0.21	1.67
CARICOM	0.64	0.53	0.65	0.60	-0.06	0.61	-2.24	-2.81	0.66
EAC	0.87	0.86	0.46	0.80	0.68	-0.58	0.85	0.42	-5.55
EU	-2.05	1.84	0.76	2.45	2.83	-5.72	5.52	1.51	-6.33
MERCOSUR	1.42	1.27	1.45	-0.04	-0.16	1.32	-4.99	-3.63	-0.82
SACU	0.64	-2.00	0.74	0.62	0.56	0.62	-3.71	0.73	0.73
LICH	-5.42	0.63	0.40	0.52	0.63	0.58	0.33	0.45	0.36

across these nine groups without accounting for the GPS and accounting for the GPS. Figure ?? shows a histogram of the t-statistics of an equality-of-means test for all covariates. Table 3 lists these statistics for each covariate except for the 115 sector fixed effects. A clear lack of balance is evident: Only 33% (24%) of the t-statistics are below 2.58 (1.96) indicating the 1% (5%) level of significance. In order to account for the GPS, we divide the GPS into 40 strata and compare the means across groups within these strata. All three methodologies used to estimate the GPS improve the balancing of covariates significantly with the unconditional density providing the best balancing property. Therefore, we focus on this specification in the text and relegate the reader to the appendix for the other specifications. The balancing of covariates improves enormously, with 95% (91%) of the t-statistics indicating equal means once we account for the GPS obtained by the unconditional density estimation. Tables 3 and 4 underline the size of improvement for all covariates. Almost all unbalanced covariates are dummy variables often concentrated in one or two groups only, making it technically impossible to achieve balance across groups even with conditioning on the GPS. For example all observations with the importer being either Switzerland or Liechtenstein are in group 1 explaining the ‘unbalanced’ t-statistic for the respective customs union dummy. The outliers in figure 3 correspond to sector dummies for *production, processing and preserving of meat, Service activities related to printing and manufacture of macaroni, noodles, couscous and similar farinaceous products* being mainly concentrated in two or three groups. To additionally tackle the imbalances, we follow Imai and van Dyk (2004) who suggest to control for the unbalanced variables when estimating the reduced form of (11).

To isolate the bilateral trade cost component, we plug the estimate of \hat{d}_{ij}^s obtained in the fixed effects regression (12) in the reduced form (11).²³ We estimate this modified reduced form using the actually observed policy treatment

$$\hat{d}_{ij}^s = k(m_{ij}^s, r(m_{ij}^s, q_{ij}^s)) + \gamma' u_{ij}^s + \xi_{ij}^s \quad (16)$$

by a polynomial of degree 3 with various interaction terms.²⁴ To account for the imbalances identified before, we follow Imai and van Dyk (2004) and estimate (16) using sector and customs union dummies as well as f_j^s as additional controls. Table 5 reports two different specifications of the estimation. Since including 115 sector dummies might absorb a lot of variation, we conduct the estimation with 115 sector dummies corresponding to the 4-digit ISIC Rev. 3 classification

²³Notice that a direct estimation of (11) with importer-sector specific effects would not allow us to identify the importer-sector specific impact of nontariff policy barriers.

²⁴The choice of degree is based on an Akaike information criterion. Robustness checks using multivariate adaptive regression splines have been conducted and can be obtained on request.

Table 5: Estimation of reduced form (16).

τ_{ij}^s	3.16***	(6.85)	0.24	(0.39)
η_j^s	-2.04***	(-4.76)	-8.04***	(-14.04)
$(\tau_{ij}^s)^2$	-1.92*	(-1.87)	1.22	(0.81)
$(\eta_j^s)^2$	3.77***	(3.36)	18.09***	(12.35)
$(\tau_{ij}^s)^3$	0.19	(0.31)	-1.46	(-1.44)
$(\eta_j^s)^3$	-2.24**	(-2.36)	-10.07***	(-8.45)
$\tau_{ij}^s \times GPS$	-0.09*	(-1.92)	0.04	(0.74)
$\eta_j^s \times GPS$	0.12***	(3.41)	0.39***	(8.91)
$\tau_{ij}^s \times \eta_j^s$	-0.42	(-0.34)	8.84***	(6.00)
$(\tau_{ij}^s \times GPS)^2$	0.04**	(2.18)	0.05**	(2.50)
$(\eta_j^s \times GPS)^2$	-0.01*	(-1.90)	-0.05***	(-5.88)
$(\tau_{ij}^s \times \eta_j^s)^2$	5.62	(0.76)	0.20	(0.02)
$(\tau_{ij}^s \times GPS)^3$	-0.00**	(-1.98)	-0.01***	(-4.03)
$(\eta_j^s \times GPS)^3$	0.00	(1.37)	0.00***	(5.75)
$(\tau_{ij}^s \times \eta_j^s)^3$	-20.98*	(-1.85)	-39.79***	(-2.93)
<i>GPS</i>	-0.05***	(-4.13)	-0.12***	(-8.23)
<i>GPS</i> ²	0.00***	(3.90)	0.01***	(5.22)
<i>GPS</i> ³	-0.00***	(-3.70)	-0.00***	(-4.66)
Contiguity	-0.45***	(-11.12)	-0.42***	(-9.52)
Common lang. (off.)	0.14***	(3.93)	0.17***	(3.89)
Common lang. (ethno.)	-0.47***	(-14.42)	-0.48***	(-12.61)
Colony	-0.19***	(-4.72)	-0.24***	(-4.35)
Common colonizer	-0.31***	(-7.45)	-0.26***	(-5.69)
Current colony	-0.43	(-0.40)	-0.80	(-0.63)
Colony (1945)	-0.22***	(-3.93)	-0.24***	(-3.42)
Same country	-0.06	(-0.71)	-0.10	(-1.09)
log(Distance)	0.29***	(29.72)	0.29***	(25.53)
f_j^s	-0.06***	(-25.15)	0.04***	(20.52)
CAN	0.75***	(23.02)	1.02***	(28.54)
CARICOM	-0.12	(-0.84)	0.34**	(2.08)
EAC	-0.17**	(-1.99)	-0.72***	(-7.71)
EU	0.02	(1.00)	-0.24***	(-8.40)
MERCOSUR	0.52***	(11.92)	0.50***	(10.66)
SACU	0.80***	(7.28)	1.03***	(9.28)
LICH	0.67*	(1.87)	1.08***	(3.10)
Sector fixed effect	Yes (4-digit level)		Yes (3-digit level)	
Customs union fixed effect	Yes		Yes	
<i>N</i>	32664		32664	

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

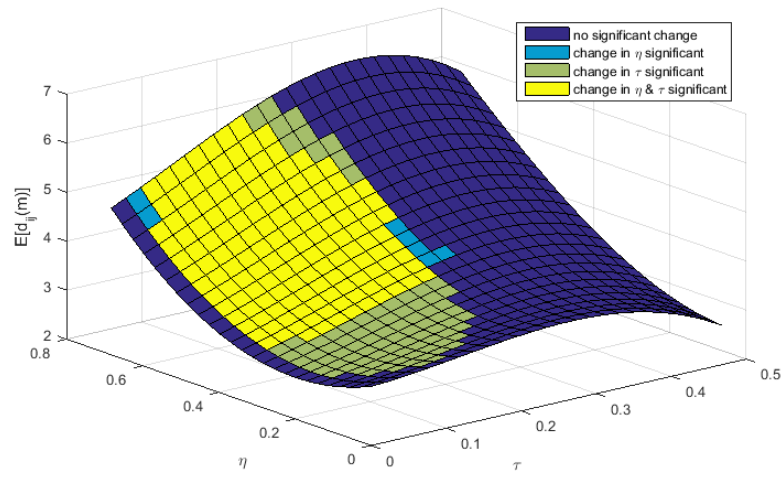
and with 55 sector dummies corresponding to the 3-digit industry classification. While the coefficients have no causal interpretation, the significant coefficients on the GPS variables and on variables interacted with the GPS indicate its importance in accounting for the endogeneity bias.

With an estimate of the reduced form at hand we are able to calculate the average causal effect of trade policy on trade costs for any level of potential policy variables. Taking the 1st and 99th percentile of each policy variable as boundaries we construct an even grid of 25×25 potential policy vectors m and estimate $\hat{\mathcal{K}}(m) = E \left[k \left(m, r(m, q_{ij}^s) \right) \right]$ by

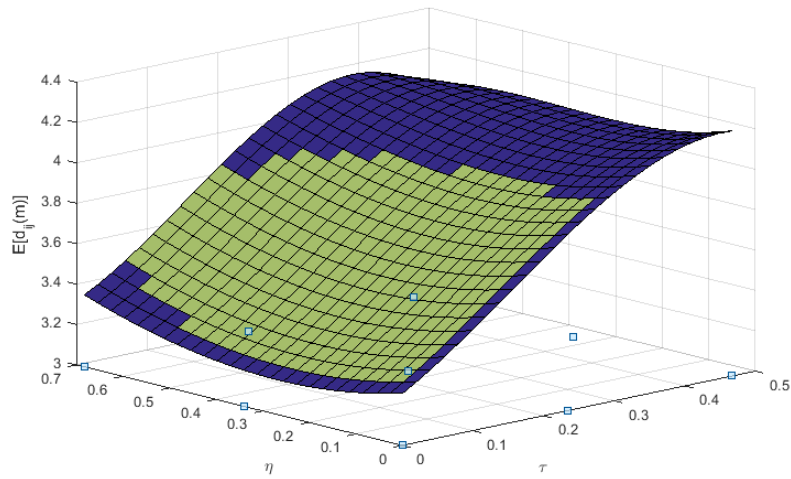
$$\hat{\mathcal{K}}(m) = \frac{1}{n} \sum_{i=1}^J \sum_{j=1}^J \hat{k}(m, \hat{r}(m, q_{ij}^s)). \quad (17)$$

Figure 4 plots the dose response function (17) for the specification with 4-digit and 3-digit

Figure 4: Average dose response function plotted for a 25 grid of potential policy vectors.



(a) Based on reduced form with 3-digit sector dummies.



(b) Based on reduced form with 4-digit sector dummies.

sector dummies, respectively. The (10 %) significance levels indicated in the graphs refer to a change in trade costs implied by a one-grid-point change of a potential policy vector and are based on 100 bootstrap samples. Interestingly, the effects of tariff changes are very robust to the inclusion of sector dummies, while a lot of the effects of nontariff barriers is absorbed by these dummies since nontariff barriers are only varying on the sector-importer dimension. Two results stand out: First, the effects of tariff changes are diminishing and eventually vanishing for high tariffs. Second, the effect of tariff changes is very much interrelated with nontariff barriers: The range of tariffs where a change in tariffs has a significant impact on bilateral trade costs is substantially smaller for those tariffs that come with high nontariff barriers. In general, the main heterogeneity of both tariff and nontariff effect comes from the concentration of effects in the area of low to medium tariffs. The effect of nontariff barriers is mostly concentrated in the area of high nontariff barriers where the effects of tariff-barriers are restricted to a quite small set of tariffs.

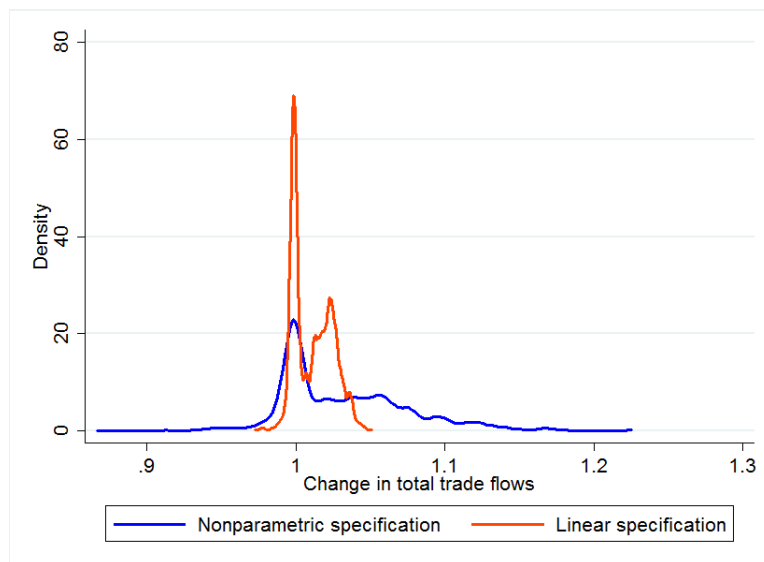
4 General equilibrium

A given change in bilateral trade costs has differential effects on trade flows due to general equilibrium effects that have not been considered in the previous analysis. In most papers on international trade these general (and partial) equilibrium effects are the only way trade cost change can have heterogeneous effect across countries. The impact of tariffs on trade costs itself, however, is taken as a one-to-one effect. Therefore, we take the general equilibrium model of section 2 and analyse the effects assuming a heterogeneous impact of tariffs on trade costs versus the classical one-to-one relationship. Specifically, we take equation (8) and conduct the comparative static analysis of a change in trade costs. Since the general equilibrium analysis requires a full matrix of trade flows including domestic absorption, we are restricted by data requirements to a subset of 36 countries + a residual country named Rest of the World (ROW) that is taken as the numéraire country and 116 sectors.²⁵

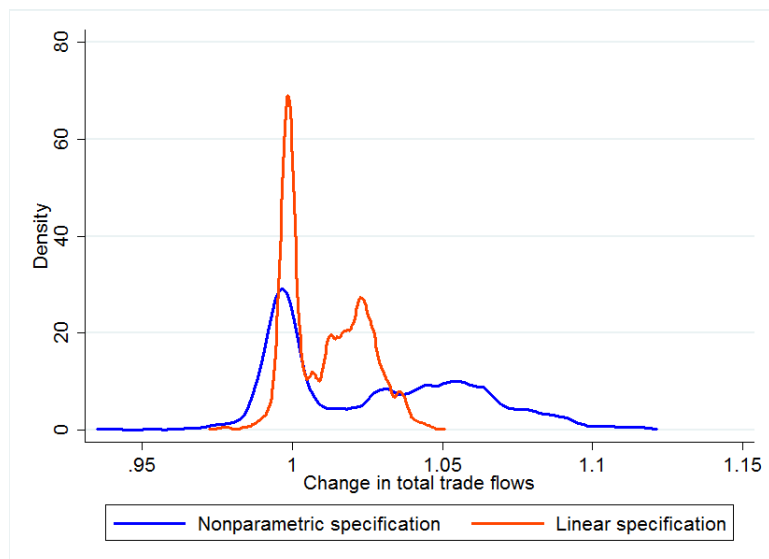
The experiment we consider is a decrease in the tariff rate by one grid point, hence a decrease in the tariff by 2%. In the classical linear case, this will lead to a change in bilateral trade costs of $\hat{D}_{linear} = e^{-0.02}$. The heterogeneous effects of our specification are accounted for by matching the true tariff and nontariff barriers of the respective country-sector pair to the closest point of the 25×25 grid to define $\log(D_{het})$ and $\log(D'_{het})$, respectively. We normalize the change

²⁵The countries considered in the general equilibrium exercise are AUS, AUT, BEL, BGD, CAN, CHN, CZE, DEU, DNK, ESP, FIN, FRA, GBR, GRC, IDN, IND, ITA, JPN, KOR, LTU, LVA, MEX, MYS, NLD, NOR, NZL, OMN, PHL, POL, PRT, ROW, SGP, SVN, SWE, THA, TUR, USA.

Figure 5: Density plot of trade flow changes in general equilibrium.



(a) Based on reduced form with 3-digit sector dummies.



(b) Based on reduced form with 4-digit sector dummies.

Table 6: Change in exports (unweighted mean across importers) in general equilibrium by country across specifications.

Exporter	Linear specification		Nonpar. specification (3-digit sector dummies)		Nonpar. specification (4-digit sector dummies)	
	All	Treated	All	Treated	All	Treated
AUS	1.0207	1.0245	1.0542	1.0658	1.0498	1.0592
AUT	1.0074	1.0212	1.0208	1.0564	1.0179	1.0510
BEL	1.0083	1.0225	1.0273	1.0724	1.0206	1.0560
BGD	1.0014	1.0138	0.9968	1.0412	1.0029	1.0347
CAN	1.0205	1.0248	1.0559	1.0677	1.0482	1.0586
CHN	1.0191	1.0217	1.0589	1.0676	1.0467	1.0534
CZE	1.0081	1.0220	1.0244	1.0629	1.0206	1.0556
DEU	1.0069	1.0205	1.0202	1.0570	1.0168	1.0501
DNK	1.0083	1.0229	1.0235	1.0642	1.0202	1.0560
ESP	1.0099	1.0256	1.0306	1.0756	1.0248	1.0641
FIN	1.0077	1.0206	1.0217	1.0572	1.0187	1.0500
FRA	1.0072	1.0198	1.0202	1.0536	1.0173	1.0480
GBR	1.0074	1.0207	1.0205	1.0555	1.0177	1.0499
GRC	1.0112	1.0275	1.0376	1.0906	1.0278	1.0684
IDN	1.0141	1.0176	1.0456	1.0498	1.0328	1.0398
IND	1.0237	1.0274	1.0797	1.0930	1.0592	1.0687
ITA	1.0085	1.0238	1.0265	1.0702	1.0211	1.0589
JPN	1.0159	1.0181	1.0402	1.0460	1.0388	1.0444
KOR	1.0164	1.0186	1.0360	1.0416	1.0402	1.0460
LTU	1.0099	1.0267	1.0329	1.0832	1.0251	1.0671
LVA	1.0089	1.0205	1.0140	1.0436	1.0192	1.0479
MEX	1.0119	1.0220	1.0393	1.0683	1.0283	1.0531
MYS	1.0079	1.0112	1.0291	1.0401	1.0191	1.0271
NLD	1.0074	1.0196	1.0226	1.0580	1.0182	1.0481
NOR	1.0083	1.0163	1.0272	1.0463	1.0200	1.0395
NZL	1.0162	1.0191	1.0321	1.0401	1.0362	1.0437
OMN	1.0165	1.0196	1.0492	1.0610	1.0395	1.0478
PHL	1.0115	1.0157	1.0462	1.0541	1.0273	1.0368
POL	1.0096	1.0249	1.0293	1.0757	1.0239	1.0627
PRT	1.0082	1.0224	1.0265	1.0722	1.0202	1.0552
ROW	0.9959	.	0.9913	.	0.9904	.
SGP	1.0121	1.0159	1.0283	1.0380	1.0287	1.0376
SVN	1.0081	1.0226	1.0256	1.0685	1.0201	1.0558
SWE	1.0065	1.0195	1.0193	1.0550	1.0161	1.0481
THA	1.0127	1.0163	1.0365	1.0457	1.0309	1.0394
TUR	1.0114	1.0249	1.0365	1.0766	1.0284	1.0623
USA	1.0192	1.0235	1.0543	1.0663	1.0451	1.0555
Total	1.0109	1.0209	1.0319	1.0597	1.0264	1.0508

in trade costs to one for all bilateral trade flows subject to a zero tariff and those countries where the closest grid point found is on the boundaries of the grid. Figure 5 plots a density plot of the bilateral trade flow changes induced by this experiment for the linear as well as the nonparametric specification. In both cases, the effect of a reduction of tariffs is more spread as compared to the linear specification. Notice that the peak around unity is mainly induced by the country-sector pairs without any change in trade costs. At first glance, the heterogeneity of the effects is visible by the smoothing-out of the peak that can be seen in the linear specification. To get a more detailed picture of the distribution of effects, table 6 reports the first moments of the distribution of trade flow changes for each exporter.²⁶ Clearly, the effects are substantially higher for the nonparametric specifications: If we look at treated trading pairs only (hence, pairs where the change in trade costs is not unity) the mean increase in trade flows is 6% in the nonparametric specification as compared to only 2% in the linear specification.

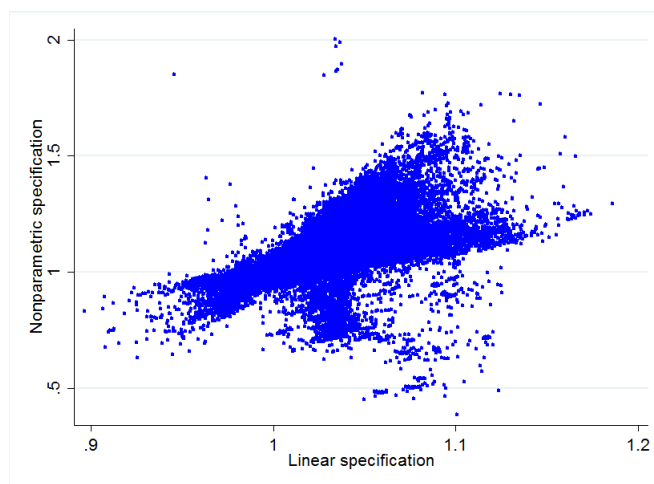
In addition to the effect on the overall distribution of general equilibrium effects, an additional source of heterogeneity is the within-variation of trade flow pairs across specifications. How does the same country-sector pair behave in different specifications? Figure 6 (a) plots the predicted change in sectoral trade flows in the linear specification against the outcome of the nonparametric specification (3-digit sector dummies). While there is a certain degree of correlation (43%), there is also strong evidence for within-variation, hence country-sector pairs where the predictions are very sensitive to the linear restriction. Figures 6 (b) and (c) shed light on the source of this heterogeneity. For pairs with very high tariffs (above the 95th percentile) there is basically no correlation between the linear prediction and the nonparametric prediction. High nontariff barriers, in turn, seem to lead to a much stronger effect of a respective change in tariffs as compared to the average effect.

5 Conclusion

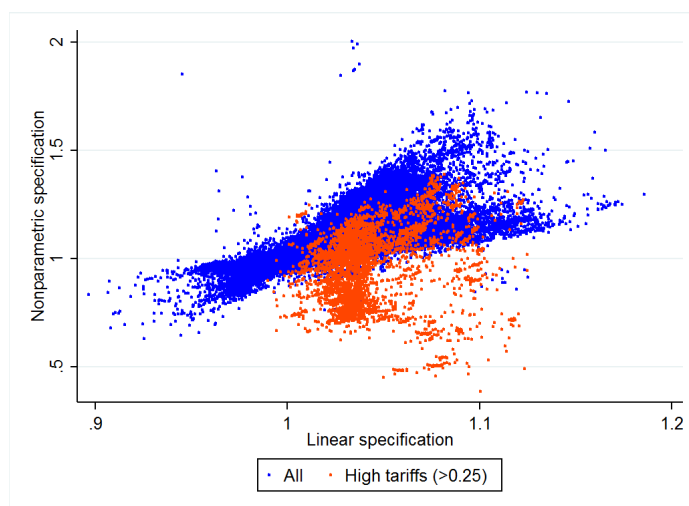
This paper tries to account for the potential heterogeneous effects of tariff and nontariff policy barriers on trade costs and on trade flows. Accounting for both potential endogeneity and nonlinearities in the impact of policy barriers on trade costs we find that the heterogeneity with respect to tariff barriers is mainly due to the lack of any effect for countries with high tariff and/or high nontariff barriers. Non-tariff barriers are especially relevant in determining trade flows for rather high levels of these barriers. The heterogeneity of the impact of policy variables on trade costs translates also to the general equilibrium outcome of changes in trade flows induced by trade cost changes. Compared to a restriction to linearity, the nonparametric

²⁶Note that the reported numbers are an unweighted mean across all importers.

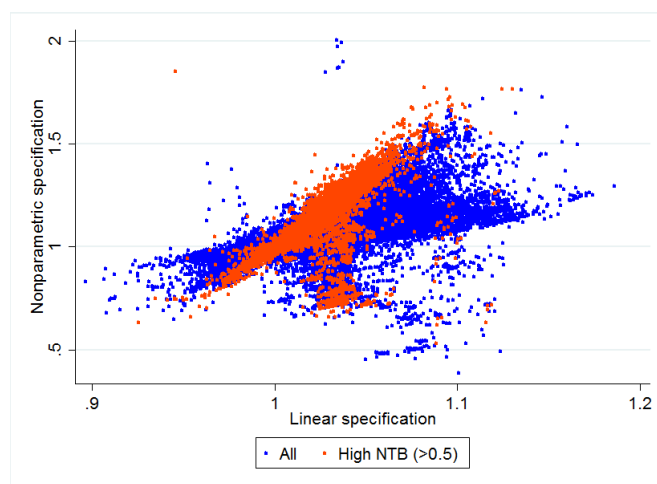
Figure 6: Within-variation: Predicted change in trade flows in the linear vs. nonparametric specification (3-digit sector dummies).



(a) Scatterplot for all trade flows (sector-level).



(b) Scatterplot with high-tariff pairs highlighted.



(c) Scatterplot with high-nontariff-barrier pairs highlighted.

specification increases the range of potential changes in trade flows substantially and rises the median and mean effect by a factor of up to three.

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Appendix

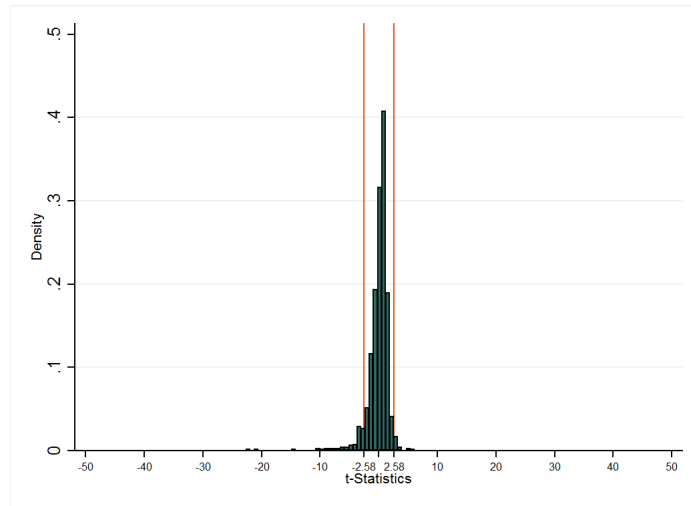
Table 7: t-statistics of equality-of-means test for non-fixed-effects covariates accounting for the GPS (normal density).

Covariates	Group								
	1	2	3	4	5	6	7	8	9
f_i^s	0.46	0.84	-0.70	-1.04	0.19	1.45	1.49	1.04	1.44
f_j^s	-1.13	0.32	-0.73	-0.33	0.77	1.75	2.87	2.04	3.52
log(Distance)	-1.02	0.75	0.96	-0.60	-0.44	0.02	0.90	0.11	0.24
Contiguity	0.33	-0.58	-0.69	-0.22	-0.50	0.47	0.03	0.31	1.00
Common lang. (off.)	-0.73	0.16	-0.83	-0.53	-0.28	-0.77	-0.99	-0.12	0.53
Common lang. (ethno.)	-0.64	-0.52	-1.01	-0.47	-0.51	-0.58	-1.28	0.13	0.52
Colony	0.08	0.24	0.08	0.62	0.12	-0.20	0.74	0.37	-0.01
Common colonizer	1.21	-0.25	-0.92	0.12	-0.00	0.49	-1.98	-1.42	-0.63
Colony (1945)	0.80	0.48	-0.10	0.57	-0.01	0.05	-0.01	-0.18	-0.44
Same country	0.28	-0.38	-0.93	0.11	-0.02	0.49	-0.39	0.10	0.66
$\bar{f}_{-i,-j}$	0.24	0.85	-0.97	-0.92	0.29	1.76	1.74	1.75	2.55
mean(Distance)	1.02	-0.75	-0.96	0.60	0.44	-0.02	-0.90	-0.11	-0.24
CAN	1.77	2.08	2.02	-2.15	-2.62	-3.18	0.50	0.37	1.65
CARICOM	0.76	0.73	1.11	0.90	-0.37	0.81	-0.90	-2.53	0.74
EAC	1.30	1.59	0.24	1.51	1.14	-0.76	1.70	-0.53	-3.25
EU	-1.72	2.26	2.92	1.88	2.22	-1.19	5.84	2.95	0.84
MERCOSUR	1.14	1.39	1.53	0.66	0.02	1.51	-0.36	-0.99	-0.54
SACU	0.59	-3.06	1.04	0.82	0.92	0.97	-3.59	0.74	0.59
LICH	-1.95	0.43	0.38	0.60	0.48	0.35	0.50	0.36	0.21

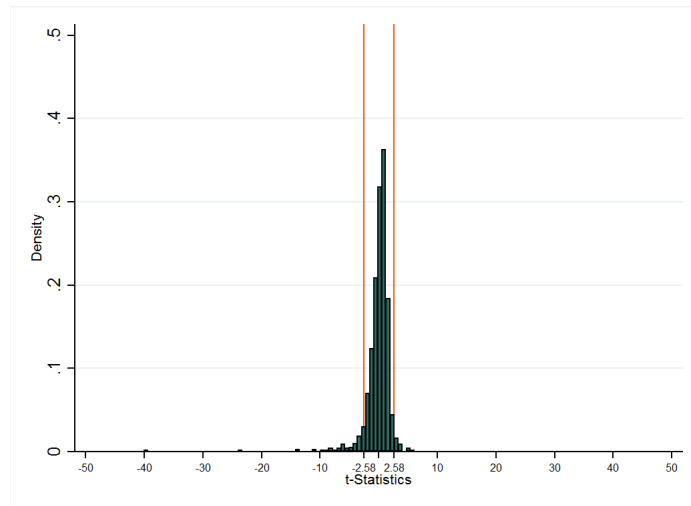
Table 8: t-statistics of equality-of-means test for non-fixed-effects covariates accounting for the GPS (log-concave density).

Covariates	Group								
	1	2	3	4	5	6	7	8	9
f_i^s	0.82	0.91	-0.26	-1.00	-0.49	1.90	1.16	1.21	-0.44
f_j^s	-0.58	0.58	0.04	-0.32	0.29	2.12	2.53	1.96	4.82
log(Distance)	-0.38	0.82	1.09	-0.38	-0.38	-0.17	0.54	0.43	1.00
Contiguity	0.08	-0.67	-0.72	-0.61	-0.71	0.46	-0.06	0.18	0.45
Common lang. (off.)	-0.71	0.04	-0.70	-0.70	-0.22	-1.03	-0.72	-0.04	-0.33
Common lang. (ethno.)	-0.57	-0.59	-0.86	-0.69	-0.52	-0.84	-1.06	0.09	0.58
Colony	0.37	0.28	0.11	0.58	0.03	-0.25	0.82	0.34	-1.44
Common colonizer	0.31	-0.34	-0.94	0.22	-0.15	0.47	-1.89	-1.31	-1.35
Colony (1945)	0.78	0.48	-0.06	0.57	-0.05	0.09	0.08	-0.31	-2.15
Same country	-0.04	-0.36	-0.71	0.03	-0.04	0.38	-0.32	-0.02	0.30
$\bar{f}_{-i,-j}$	0.52	0.99	-0.54	-0.86	-0.40	2.32	1.39	1.63	3.52
mean(Distance)	0.38	-0.82	-1.09	0.38	0.38	0.17	-0.54	-0.43	-1.00
CAN	1.67	2.02	2.18	-2.18	-2.88	-3.04	0.21	0.32	1.29
CARICOM	0.79	0.73	1.45	0.93	-1.12	1.09	-0.13	-3.15	1.07
EAC	1.28	1.52	0.15	1.50	1.07	-0.16	1.78	-0.26	-11.21
EU	-1.56	2.37	3.08	1.86	2.47	-1.25	5.92	2.56	-2.15
MERCOSUR	1.08	1.31	1.61	0.72	-0.24	1.46	-1.03	-1.36	-2.91
SACU	0.59	-3.93	1.53	0.61	0.90	1.76	-5.30	1.13	0.64
LICH	-1.52	0.44	0.45	0.54	0.40	0.41	0.49	0.66	0.23

Figure 7: Distribution of t-statistics of equality-of-means test for all covariates without accounting for the GPS.



(a) Normal density.



(b) Log-concave density.