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Heterogeneous effects of economic integration agreements☆

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ABSTRACT

It is now widely accepted that economic integration agreements (EIAs) and other trade-policy liberalizations contribute to nations' economic growth and development and help alleviate poverty. However, the economic effects of such policies vary across countries' economic structures; for instance, developing countries face higher fixed trade costs (partly due to higher government border-crossing costs and weaker port infrastructures). We offer three potential contributions. First, we extend a standard Melitz general equilibrium trade model with firm heterogeneity to show how variable-cost and fixed-cost "trade elasticities" associated with trade liberalizations are heterogeneous and endogenous to *levels* of country-pairs' bilateral policy and non-policy, variable and fixed trade costs – even allowing for CES preferences and an untruncated Pareto distribution of productivities. Using associated comparative statics, we provide several explicit predictions of the heterogeneous (variable- *and* fixed-cost) bilateral extensive-margin, intensive-margin, and trade elasticities. Second, we provide empirical support for the theoretical hypotheses. Trade elasticities vary across particular settings. Third, we demonstrate the relevance of these theoretical and empirical results for *ex ante* trade-flow predictions of potential EIAs. For instance, we show that a 10 percent lower average per capita income of a country-pair is associated with a *60 percent* higher partial EIA effect. Moreover, we show empirically that 95–99 percent of the welfare (or probability) estimates of EIA liberalizations between 1,358 North-North, North-South, and South-South country-pairs can be explained by our heterogeneous EIA partial treatment effects.

"In general, trade liberalization is an ally in the fight against poverty …. " (Trade Liberalization and Poverty: A Handbook (2001), p. 3)

*"From an empirical point of view, we would like to have substantially richer evidence on the magnitude of the trade elasticity based on trade policy variation, and **most importantly**, on the question of whether the trade elasticity appears to be invariant across time and space, or is **dependent on the particular setting**." (Goldberg and Pavcnik (2016), p. 31; bold added)*

*"The ultimate effect of a free trade area **depends on the particular characteristics** of member countries …. " (Economic Commission for Africa (2017), p. 65; bold added)*

1. Introduction and relevant literature

It is now widely accepted that economic integration agreements (EIAs) and other trade-policy liberalizations contribute to nations' eco-

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nomic growth and development. EIAs have proliferated among North-North (N-N), North-South (N-S), and South-South (S-S) country-pairs. While such agreements inevitably alter distributions of income within countries, for the most part EIAs are believed to raise economic welfare.[1]

A major recent advance in the international trade literature – in the wake of and building upon theoretical developments associated with firm heterogeneity and export fixed costs – is the development of the "new quantitative trade models," cf. Redding (2011), Arkolakis et al. (2012), and Head and Mayer (2014). These models – explored in detail in Costinot and Rodriguez-Clare (2014) – provide calculations of general equilibrium trade and welfare effects of trade liberalizations using exogenous (variable-cost) "trade elasticities" estimated from structural gravity equations combined with aggregate bilateral trade data. These "mid-sized" numerical general equilibrium models are built on sound microeconomic foundations, are transparent, and have limited data requirements. Moreover, Head and Mayer (2014) demonstrated that estimates of welfare effects of economic integration agreements (EIAs) can be computed once one has partial treatment effects from a properly specified gravity equation with EIA dummy variables and an *exogenous* trade-elasticity (parameter) value.

However, as our quote above from Goldberg and Pavcnik (2016) notes, an important unresolved (and hardly explored) issue is whether – and by what factors – trade elasticities with respect to trade-policy changes vary "across time and space," that is, are sensitive to "particular settings"; this is particularly important in contrasting trade elasticities for N-N, N-S, and S-S EIAs.[2] This is the issue we explore in this paper. We do so by addressing three particular questions. First, *how* are trade elasticities – fixed-cost-trade-policy trade elasticities as well as variable-cost ones – theoretically related to *levels* of fixed and variable trade-cost variables, which vary dramatically between N-N, N-S, and S-S pairs? Second, is there convincing *empirical* evidence supporting these theoretical interactions? Third, how important *quantitatively* is the heterogeneity in partial equilibrium trade impacts in determining the general equilibrium welfare impacts of trade-policy liberalizations?

To address these questions, this paper offers three potential contributions. First, extending a standard Melitz model of trade, we show theoretically the heterogeneity of trade elasticities to given *ad valorem* tariff-rate and to export-fixed-cost-policy changes depending upon *levels* of variable and export-fixed costs. We show theoretically how extensive-margin, intensive-margin, and trade elasticities are endogenous to the levels of theoretical bilateral variable and fixed, policy and non-policy trade costs – even with CES preferences and with an untruncated Pareto productivity distribution. Ours is not the first paper to address theoretically the endogeneity of the (variable-cost) trade elasticity. Melitz and Redding (2015) note that the exogeneity of this trade elasticity stems from the typical assumption of an untruncated Pareto distribution for firms' productivities; they show that assuming a *truncated* Pareto distribution endogenizes the trade elasticity. Melitz and Redding (2015) also note an emerging empirical literature on heterogeneous (or endogenous) trade elasticities. For instance, Helpman et al. (2008), or HMR, find empirical evidence for endogenous elasticities of trade with respect to distance in the context of a truncated Pareto distribution of productivities. They showed that, when trade costs related to distance fall, the response of the extensive trade margin is considerably larger for developing countries than for developed countries. To anticipate just one of our results, we find evidence that the response of the extensive trade margin to an EIA is *also* considerably larger for *developing* countries. Moreover, Novy (2013), in a model with homogeneous firms, finds that exogenous trade elasticities are a feature of models with constant-elasticity-of-substitution (CES) preferences; using transcendental logarithmic (translog) preferences, Novy (2013) demonstrates that the trade elasticity can be endogenous. Our theoretical model is distinct from Melitz and Redding (2015) and Novy (2013) by finding theoretically endogenous trade elasticities with respect to trade-policy changes in the context of two assumptions common to the new quantitative trade models: CES preferences and an untruncated Pareto productivity distribution. The basic rationale is that, by introducing additively separable components of variable trade costs and of fixed trade costs, the trade elasticities with respect to a single component become variable. Moreover, our model – incorporating network effects as in Krautheim (2012) – is likely the first paper to address the endogeneity of trade elasticities to *fixed-export-cost* changes, an issue suggested in Goldberg and Pavcnik (2016) and Limao (2016).

Second, we evaluate empirically our theoretical hypotheses. We provide empirical evidence confirming our theory and demonstrate the heterogeneity of EIAs' trade effects depending upon country-pairs' geographic, cultural, institutional, and development characteristics. Extending here Baier et al. (2014) and Head and Mayer (2014), this is the first paper to show evidence that extensive-margin, intensive-margin, and trade-flow EIA elasticities are indeed sensitive to levels of (observable) bilateral variable and fixed, policy and non-policy export costs in a manner consistent with theoretical comparative statics. Trade elasticities with respect to trade-policy changes *do* vary across "particular settings." Geographic, cultural, institutional, and development country-pair characteristics all significantly influence the extensive margin elasticity, whereas primarily geographic variables (distance and adjacency) influence the intensive margin elasticity, consistent with our theory.[3]

Third, our framework allows us to put to *ex ante* use the partial effects of EIAs. Historically, the heterogeneity of EIAs' effects on members' bilateral trade could only be evidenced *ex post* using separate EIA dummy variables for various agreements; however, such estimates are weak due to insufficient variation in the right-hand-side (RHS) dummies. By explaining the heterogeneity of EIAs' effects according to theoretically-motivated factors, one can use the heterogeneous partial (treatment) effects for *ex ante* predictions, which we motivate later, and we demonstrate empirically that the partial effect of an EIA tends to be much larger for a pair of *developing* economies. Moreover, in the context of the new quantitative trade models, we demonstrate empirically using two approaches *how sensitive quantitatively* general equilibrium welfare effects of EIA liberalizations are to the bilaterally heterogeneous (partial) trade elasticities. In one approach, we calculate the general equilibrium welfare effects for importers of 1,358 bilateral EIA liberalizations among N-N, N-S, and S-S country-pairs. Consistent with theory, we show that 98–99 percent of the variation in these 1,358 welfare changes can be explained by the variation in two statistics: the estimated bilateral EIA dummy coefficient and the share of the importer's national expenditures on exports from the EIA partner. In the other approach, based upon the methodology in Baier and Bergstrand (2004) we show that the probability of two countries having an EIA – which in

[1] Seminal empirical studies of the effect of openness and trade-policy liberalizations on economic growth and development include Esfahani (1991), Harrison (1996), Frankel and Romer (1999), Awokuse (2007), Badinger (2008), and Chang et al. (2009). For an overview of how historical factors involving international trade policies matter for economic development, see Nunn (2014).

[2] On March 21, 2018, the heads of 44 of 55 African countries signed a new Continental Free Trade Area agreement, which is considered – according to Economic Commission for Africa (2017) – the "first flagship project" and a "key initiative" of the African Union Agenda 2063 for making progress toward the UN's Sustainable Development Goals. We will conclude with how this paper's methodology could be applied to estimate better the trade and welfare gains from such an agreement.

[3] Baier et al. (2014) showed that extensive-margin changes from EIAs can be considerable once endogeneity of EIAs is accounted for properly econometrically, supporting notions raised in Trefler (1993) and Trefler (2004). Also, Baier et al. (2016) provide evidence of heterogeneous EIA trade elasticities, but do not link these to geographic, cultural, institutional, and development country-pair characteristics.

the context of their theory is related to the net welfare gain from such EIA – is highly correlated with the heterogeneous EIA coefficients and the trade shares. In fact, we show that the estimated heterogeneous EIA coefficients and bilateral trade shares – accounting also for other economic factors influencing the probability of an EIA also – can explain up to 95 percent of the variation of such probabilities, consistent with our theory.

The remainder of the paper is as follows. In section 2, we extend a standard Melitz model of trade to first motivate how the interactions of exogenous factors influencing fixed export costs – exogenous non-policy fixed export costs and exogenous policy fixed export costs – with endogenous fixed export costs associated with "network effects" (as raised in Krautheim (2012)) can explain theoretically the sensitivity to fixed export cost levels of the elasticity of the extensive margin of trade flows with respect to variable tariff rates – even with CES preferences and an untruncated Pareto productivity distribution. Second, we show that the elasticity of the extensive margin of trade flows with respect to fixed policy export costs is also sensitive to the levels of fixed export costs. Third, we show furthermore that the elasticity of the extensive margin of trade flows with respect to fixed policy export costs is sensitive to the *relative levels* of exogenous policy and non-policy fixed export costs. Fourth, incorporating more economically plausible representations of variable trade costs into the framework (as raised in Anderson and van Wincoop (2004)) yields an endogenous intensive-margin variable-tariff-rate elasticity as well.

In sections 3 and 4, we provide an empirical analysis of our theoretical hypotheses using HMR's "geographic, institutional, and cultural" variables as proxies for policy and non-policy fixed export costs in a gravity model. As noted in Nunn and Trefler (2014), good institutions are a potential source of comparative advantage; recent trade models with firm heterogeneity find evidence that good institutions can expand trade at the extensive margin. Specifically, in the baseline specifications we show that distance, adjacency, and typical gravity dummy variables reflecting common institutional and cultural country characteristics (the latter capturing exogenous policy and non-policy, respectively, fixed export costs) explain well the heterogeneity in EIA dummy variables' partial effects on the *extensive* (product) margin. Moreover, we show that distance and adjacency – influencing variable transport costs – explain well the heterogeneity in EIA dummies' partial effects on the *intensive* (product) margin. Furthermore, we show how the heterogeneous EIA effects are related to various aspects of development. To the best of our knowledge, only two studies have estimated heterogeneous EIA effects using interaction terms like here to avoid the dilemma of a multitude of individual dummies that yield econometrically weak coefficient estimates. Vicard (2011) investigated empirically interactions of numerous economic variables with EIA dummies, but the study was not guided by theory and so interaction effects lacked economic interpretation. Cheong et al. (2015) examined empirically interactions of EIA dummies only with measures of GDP size similarity and per capita income similarity and found significant effects. Like Vicard (2011) though, Cheong et al. (2015) was not guided by theory. Also, both of those studies looked only at aggregate trade flows. Our study is unique by offering theoretical guidance from a Melitz heterogeneous firms model to understand the roles of fixed and variable export costs – with or without network externalities and with an untruncated Pareto distribution – for explaining heterogeneous EIA effects, for explaining differential EIA effects – quantitatively and qualitatively – on intensive and extensive (product) margins, and for controlling for various degrees of EIA liberalization (as raised in Kohl et al. (2014)).[4] We employ the Hummels and Klenow (2005) product-margin-decomposition methodology, as in Baier et al. (2014), to explore empirically how various core gravity-equation variables influence such margins' EIA effects. This section also provides a robustness analysis of our main results to alternative nontradable goods' "cutoffs," interaction effects by type of EIA, lagged EIA effects, inclusion of other controls (in particular, aspects of development), inclusion of tariff rates, zeros in trade, and sectoral decompositions.[5]

In section 5, we show that our approach to gravity-equation modeling now makes more plausible *ex ante* use of gravity equations for predicting the partial effects of future EIAs among N-N, N-S, and S-S pairs and their likely welfare effects. First, studies such as Baier and Bergstrand (2007) and Baier et al. (2014) can help policymakers predict future partial (and then general) equilibrium effects of a planned EIA; the former (latter) study predicts the partial effect without (with) regard to type of EIA. However, those predicted partial effect estimates are homogeneous across country-pairs (based on average treatment effects). In contrast, our heterogeneous partial EIA effects are explained by several variable- and fixed-trade-cost proxies. Moreover, we show empirically that these heterogeneous partial effects are related to levels of development. Our results indicate that a 10 percent lower average per capita income of a country-pair is associated with a *60 percent* higher partial EIA effect. Second, we also will show that the heterogeneity of EIA partial effects helps to explain the likely welfare gains and predictability of EIAs. For instance, we will show that 98–99 percent of the welfare gain for country j of an EIA with country i can be explained by the heterogeneous (partial) EIA coefficient estimate along with the share of country j's expenditures on imports from country i. Put succinctly, previous gravity equations allowing for heterogeneous partial effects of EIAs on trade have been limited not just by weak estimates, but allowed only *ex post* evaluation. Our paper suggests a methodology for generating robust and precise heterogeneous partial effect estimates that can also be used potentially for *ex ante* trade and welfare analysis using the new quantitative trade models. In section 6, we demonstrate the relevance of our findings to the current trade policy debate, analyzing the partial effect of "Brexit" from the European Union (EU), as well as the potential effects of two EU members that are developing economies exiting the EU. Section 7 provides conclusions.

2. Theory

This section has four parts. In the first part, we extend a standard Melitz model of trade with heterogeneous firms, such as in Redding (2011), to incorporate additively separable tariff rates and freight rates (variable trade costs), additively separable policy and non-policy (or natural) export fixed costs, *and* additively separable exogenous and endogenous export fixed costs (or network effects, as in Krautheim (2012)). In the second part, we solve for a gravity equation. In the third part, we provide comparative statics for *ad valorem* tariff-rate changes that motivate several testable propositions. In the fourth part, we provide comparative statics for policy *fixed* export-cost changes that motivate several other testable propositions.

2.1. The model

Our theoretical model is an extension of the Redding (2011) version of the Melitz (2003) model. Our model has four economically plausible

[4] We intentionally use an untruncated Pareto distribution for productivities to distinguish the economic channels explaining our endogenous trade-cost elasticities from those channels addressed in Melitz and Redding (2015).

[5] Although we focus empirically on heterogeneous partial effects of EIA dummies, our analysis holds in principle for *ad valorem* tariff rates as well, such as in Baier and Bergstrand (2001). Our focus empirically on heterogeneous EIA dummy coefficients, rather than heterogeneous tariff-rate elasticities, is due to the "paucity" of high quality *ad valorem* tariff-rate (and nontariff-rate) data and the empirical prominence of EIA dummies in the literature, cf., Goldberg and Pavcnik (2016). EIA dummies can capture the effects of both tariff-rate *and* non-tariff-measures changes. Nevertheless, in a robustness analysis constrained by sample size due to available data, we add tariff rates.

features that distinguish it from previous Melitz models. The first is that we separate additively the gross bilateral *ad valorem* tariff rate from the *ad valorem* "freight rate," the two standard components of *ad valorem* variable trade costs in this class of models. This follows from the formulation of variable trade costs recommended in Anderson and van Wincoop (2004) and will help motivate later our empirical finding that the intensive-margin elasticity of bilateral EIAs decreases (increases) with larger bilateral distances between countries (adjacency of countries). The second is to additively separate exogenous *policy* export fixed costs from exogenous *non-policy* (or natural) export fixed costs. This feature will be important for explaining later our empirical finding that the effects of lower trade-policy-related export fixed costs (such as from forming an EIA) on bilateral extensive margins and trade flows are positively related to the presence of country-pairs' common cultural backgrounds (i.e., lower exogenous non-policy export fixed cost levels), but *negatively* related to the presence of country-pairs' common institutional backgrounds (i.e., lower exogenous policy export fixed cost levels). The third is to introduce additively separable exogenous and endogenous export fixed costs. Chaney (2008) and Redding (2011) include only exogenous export fixed costs; Krautheim (2012) includes only endogenous export fixed costs. Although Krautheim (2012) introduced endogenous export fixed costs, it was at the expense of exogenous fixed costs for the "great advantage" of solving for closed form solutions. Our model includes both exogenous and endogenous export fixed costs in an economically plausible way (additively separable), and generates endogenous tariff-rate and policy-fixed-export-cost elasticities.[6] The fourth distinguishing feature is that the additively separable exogenous and endogenous fixed costs are introduced into a Melitz model with free entry and exit, labor-market clearing, and endogenous number of varieties; the model in Krautheim (2012) did not have free entry and exit, labor-market clearing, and an endogenous number of varieties. This is not a trivial extension; accordingly, Online Appendix 1 develops this extension in a closed-economy Melitz model to prove first the existence, uniqueness, and stability of extending the Melitz model to include additively separable exogenous and endogenous fixed costs in the simplest theoretical setting possible. Online Appendix 2 develops this extension in the more general open-economy case with N countries.[7]

We assume a world economy with N countries and let L_j denote the exogenous (internationally immobile) population and labor force in country j. We assume a single industry with heterogeneous firms each producing a single differentiated product under increasing returns to scale and monopolistic competition.[8]

Consumers (workers) are identical and have the constant elasticity of substitution (CES) utility function:

$$U_j = \left(\int_{\Omega \in \Omega_j} q(\Omega)^{\frac{\sigma-1}{\sigma}} d\Omega \right)^{\frac{\sigma}{\sigma-1}} \tag{1}$$

where $q(\omega)$ denotes the quantity consumed of product ω from the set of varieties Ω_j available and σ is the elasticity of substitution in consumption across varieties ($\sigma > 1$). Consumers maximize utility subject to a standard income constraint yielding a demand function in country j for variety ω imported from country i:

$$q_{ij}(\Omega) = \left(\frac{p_{ij}(\Omega)}{P_j} \right)^{-\sigma} \left(\frac{E_j}{P_j} \right) \tag{2}$$

where $P_j = [\int_{\Omega \in \Omega_j} p(\Omega)^{1-\sigma} d\Omega]^{\frac{1}{1-\sigma}}$ and E_j is aggregate expenditure (which is equal to aggregate income in country j (Y_j) and any tariff revenue (T_j) introduced later).

Firms are assumed to have heterogeneous productivities. Entry into a market by a firm requires an exogenous cost f_i^e in country i. In order to sell in a market j, a firm has to pay a fixed cost, f_{ij}.[9] We assume that the costs (c) for a firm with productivity φ in origin i to sell q_{ij} units of output in destination j facing (gross) *ad valorem* iceberg variable trade costs τ_{ij} (hence, assuming $\tau_{ij} \geq 1$) is given by:

$$c(q_{ij}) = \frac{w_i q_{ij} \tau_{ij}}{\varphi} + w_j f_{ij} \tag{3}$$

Facing demand curve equation (2), the price charged in j by a firm in i is given by:

$$p_{ij}(\varphi) = \frac{w_i \tau_{ij}}{\rho \varphi} \tag{4}$$

where $\rho = (\sigma - 1)/\sigma$.

Up to now, our model is standard. We now introduce our first distinguishing feature. Following Anderson and van Wincoop (2004), we assume that gross tariff rates (t_{ij}) and freight rates (fr_{ij}) enter τ_{ij} additively. Anderson and van Wincoop (2004, p. 715) is the most prominent study to suggest this formulation of the trade-cost factor, $\tau_{ij} = t_{ij} + fr_{ij}$. As the U.S. Customs and Border Protection web site notes, duties are not assessed on cost-insurance-freight (CIF) charges, but rather on free-on-board (FOB) charges. Hence, for a good exported from country i to country j facing a (gross) tariff rate t_{ij}, the price at the destination (p_{ij}) should be $p_{ij} = p_i t_{ij}$ before freight costs (where p_i is the FOB price). Following Hummels (1999), Hummels and Skiba (2004), and Hummels (2007), freight costs per unit of the good ($freight_{ij}$) drive a wedge between origin and destination prices; hence, $p_{ij} = p_i t_{ij} + freight_{ij} = p_i(t_{ij} + fr_{ij})$, where $fr_{ij} = freight_{ij}/p_i$. Thus, *ad valorem* iceberg variable trade costs τ_{ij} are additively separable between an *ad valorem* gross tariff rate, $t_{ij} > 1$, and an *ad valorem* freight rate, $fr_{ij} > 0$:

$$p_{ij}(\varphi) = \frac{w_i \tau_{ij}}{\rho \varphi} = \frac{w_i (t_{ij} + fr_{ij})}{\rho \varphi}. \tag{5}$$

For simplicity in this section, we will often use τ_{ij} rather than $t_{ij} + fr_{ij}$ when the distinction between them is unnecessary; their distinction becomes more relevant in section 2.3 below.[10]

We now introduce the second and third distinguishing features of our model. We assume that fixed costs are determined by two exogenous components (A_{ij}^N and A_{ij}^P) and an endogenous component reflecting network effects ($M_{ij}^{-\eta}$). As in Krautheim (2012), we assume that the fixed costs of selling a product from i to j are inversely related to the

[6] The rationale for the third distinguishing feature was discussed above. In his final section 4, Krautheim (2012) notes, "It is quite likely, however, that in reality *some fixed costs* are entirely (or at least mainly) *independent* of the number of exporters" (p. 33; italics added). These "independent" (exogenous) fixed costs may influence the elasticity of export fixed costs with respect to the number of exporters. In fact, Krautheim (2012) concludes the last substantive section of his paper suggesting "future empirical work" should investigate the variability of trade elasticities to changes in such *exogenous* (spillover-insensitive) export fixed cost determinants, as we pursue here.

[7] Yet, a sufficient condition in our model for existence, uniqueness, and stability of the zero-profit cutoff productivity is analogous to the condition for stability in Krautheim (2012).

[8] We could introduce another (outside) homogeneous good that is traded costlessly under perfect competition to allow us to have common national wage rates, set equal to unity. However, in the last section of the paper, we want to contrast general equilibrium welfare effects with partial effects; hence, we allow national wage rates to differ, determined in the model by multilateral trade-balance constraints.

[9] As in Redding (2011), we use the term fixed costs, usually without distinction between domestic versus export. Subscripts *ii* versus *ij* distinguish between domestic versus export fixed costs. However, in some contexts where the origin and destination markets are different countries, we may use the term export fixed costs.

[10] There is just now emerging a literature on the formulation of transport costs versus tariff rates in Melitz-Chaney type models, cf. Costinot and Rodriguez–Clare (2014), Besedes and Cole (2017) and Caliendo et al. (2015, especially Appendix A). Also, we address endogeneity of fr_{ij} to φ (via $p(\varphi)$) in Online Appendix 2 and its theoretical supplement.

mass of firms in i selling in j, M_{ij}, which itself is endogenous to the model. Fixed costs are assumed to be:

$$w_j f_{ij} = w_j(A_{ij}^N + A_{ij}^P + M_{ij}^{-\eta}) \tag{6}$$

where η is the elasticity of fixed costs with respect to the mass of firms in i selling to j (and, as in Krautheim (2012), assume $0 < \eta < 1$) and we assume as is common that fixed costs of i's producers are borne in the destination country.[11]

Finally, our model departs from Krautheim (2012), both because Krautheim (2012) is a Chaney (2008) type model with an exogenous number of varieties (i.e., no free entry and exit and no labor-market clearing); this is the fourth distinguishing feature of our model. In this setting, the profits of firm φ in i selling to j (π_{ij}) are:

$$\pi_{ij}(\varphi) = \text{Max}\left[0, \left(\frac{w_i \tau_{ij}}{\rho\varphi P_j}\right)^{1-\sigma} \frac{E_j}{\sigma} - w_j(A_{ij}^N + A_{ij}^P + M_{ij}^{-\eta})\right] \tag{7}$$

Firms in i will choose to sell to j as long as profits are positive. The marginal exporter from i to j, where profits approach zero, defines the "cutoff" productivity (φ_{ij}^*):

$$\left(\frac{w_i \tau_{ij}}{\rho P_j}\right)^{1-\sigma} \frac{E_j}{\sigma}(\varphi_{ij}^*)^{\sigma-1} = w_j(A_{ij}^N + A_{ij}^P + M_{ij}^{-\eta}) \tag{8}$$

where the LHS of equation (8) is variable profits and the RHS is fixed costs. In Krautheim (2012), without the additive exogenous fixed costs $A_{ij}^N + A_{ij}^P$, one can easily solve for the cutoff productivity φ_{ij}^* (once the function for M_{ij} is specified). However, the presence of the additive factor $A_{ij}^N + A_{ij}^P$ makes the determination here of φ_{ij}^* more complex. As noted earlier, because of this complexity, we solved first for a closed-economy version of this Melitz model. This model is described in Online Appendix 1, which also provides a proof of a sufficient condition to ensure existence, uniqueness, and stability of the equilibrium values of the cutoff productivity (φ^*) and average firm profits ($\overline{\pi}$).[12]

However, equation (8) provides only an implicit solution for the zero-profit-cutoff (ZPC) productivity φ_{ij}^* (because, as we will see, M_{ij} is a function of φ_{ij}^*). Although we cannot solve explicitly for φ_{ij}^*, we show the conditions for existence of a unique and stable cutoff productivity for sales from origin i to destination j using a fixed-point argument, as in Redding (2011).[13] It will be useful to assume a distribution for firms' heterogeneous productivities. As emphasized in the introduction, we assume an untruncated Pareto distribution. The probability density function (pdf) of the productivity distribution is then $g(\varphi) = \gamma\varphi^{-(\gamma+1)}$ and the cumulative distribution function is $G(\varphi) = 1 - \varphi^{-\gamma}$, where we assume $\phi_{\min} = 1$ for convenience. Hence, $1 - G(\varphi) = \varphi^{-\gamma}$.

Given the Pareto distribution, it will be useful to make a conjecture about the functional form for M_{ij}. We conjecture that:

$$M_{ij} = \alpha_i L_i (\varphi_{ij}^*)^{-\gamma} \tag{9}$$

where α_i is solved for in Online Appendix 2. We prove this conjecture is correct in Online Appendix 2. The complete set of solutions for this Melitz model with additively separable exogenous and endogenous fixed costs is provided in Online Appendix 2.

2.2. Gravity equation

Following Redding (2011), the trade flow from country i to country j can be expressed in terms of an extensive margin and an average exports (conditional upon exporting) margin:

$$X_{ij} = \underbrace{\left[\frac{1 - G(\varphi_{ij}^*)}{1 - G(\varphi_{ii}^*)}\right] M_i}_{\text{Extensive}} \int_{\varphi_{ij}^*}^{\infty} \left(\frac{w_i \tau_{ij}}{\rho\varphi P_j}\right)^{1-\sigma} E_j \frac{g(\varphi)}{1 - G(\varphi_{ij}^*)} d\varphi \tag{10}$$

Using the Pareto distribution, Online Appendix 2 shows that equation (10) can be rewritten as:

$$X_{ij} = (\alpha_i L_i)(\varphi_{ij}^*)^{-\gamma} \left(\frac{\sigma\gamma}{\gamma - (\sigma - 1)}\right) w_j A_{ij} \left[1 + \frac{(\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\eta\gamma}}{A_{ij}}\right]. \tag{11}$$

where $\alpha_i = \left(\frac{\sigma-1}{\gamma\sigma f_i^e}\right)\left(1 - \frac{\frac{\gamma}{\sigma-1}-1}{1+\gamma}\frac{T_i}{w_i L_i}\right)$.

Equation (11) is the analogue to equation (15) in Redding (2011), where for simplicity $A_{ij} \equiv A_{ij}^N + A_{ij}^P$ (and some notation differences exist).[14] The product of the first two RHS terms capture the "extensive" margin and the product of the next three RHS terms is referred to in Redding (2011) as the "intensive" margin, though more accurately termed the "average exports (per firm)" margin, cf., Head and Mayer (2014). The average exports margin includes both the intensive margin and a "composition" margin, as Head and Mayer (2014) clarify. In Redding (2011), without endogenous fixed costs, the Pareto distribution ensures the average export margin is $\left(\frac{\sigma\gamma}{\gamma-(\sigma-1)}\right) w_j A_{ij}$. In our case with endogenous fixed export costs, we have an extra term, the last RHS term in brackets above, with two implications. First, as in Krautheim (2012), a one percent fall in the *ad valorem* tariff rate would reduce φ_{ij}^* by more than one percent (and increase trade by more than γ percent), because of lower export fixed costs $(\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\eta\gamma}$ (which is the magnification effect). Second, in our framework with independent exogenous export fixed costs, the magnification effect is sensitive to the *level* of exogenous export fixed costs A_{ij}; the lower is A_{ij}, the higher is the magnification effect. This is the intuition behind the *endogenous* trade elasticities associated with the endogenous fixed export costs discussed in the next two sections of comparative statics below, as well as a rationale for introducing EIA dummy variable *interaction terms* later in empirical specifications.

2.3. Comparative statics for ad valorem tariff rates

In this section and the next, we use partial equilibrium comparative statics to illustrate several novel insights. We choose to examine partial comparative statics in this section since our econometric exercise (in section 3) is intended only to shed light on heterogeneous and endogenous *partial* effects of EIA formations and enlargements, holding income changes (w_j) constant. General equilibrium effects (allowing for w_j changes) will be addressed in section 5.[15]

For tractability, in this section we examine three comparative statics from the model; comparative statics 1–3 are related to an exogenous

[11] We discuss later in section 3 how the exogenous component determining natural fixed export costs, A_{ij}^N, is likely influenced by (observable) geographic and cultural factors such as bilateral distance and the presence or absence of common land borders, official languages, and predominant religions. By contrast, the level of policy-oriented fixed export costs, A_{ij}^P, is likely influenced by (observable) institutional similarities such as common legal origins and colonial histories. See Costinot and Rodriguez-Clare (2014), p. 212 on the common assumption regarding payment of fixed export costs in the importing country. Finally, we can assume, with no loss of generality, that the endogenous network spillover only applies to exporters, i.e., international trade. While such an assumption is unnecessary for the results in sections 2–4, this assumption will ensure in section 5 that welfare can be measured using the standard two sufficient statistics discussed in Arkolakis et al. (2012).

[12] It will turn out that this condition is analogous to one assumed in Krautheim (2012) to ensure an interior solution.

[13] See Online Appendix 2.

[14] In the case of zero rebated tariff revenue, the expression is identical to Redding (2011). Also, most empirical estimates of γ and σ imply $\frac{\frac{\gamma}{\sigma-1}-1}{1+\gamma}$ is a small fraction.

[15] Note that by assuming in our model an untruncated Pareto distribution, the effects here complement those addressed using a truncated Pareto distribution in Melitz and Redding (2015) and in HMR.

change in *ad valorem* bilateral tariff rates ($d\ln t_{ij}$). In section 2.4 later, we solve for three comparative statics related to an exogenous change in policy-oriented bilateral export fixed costs ($d\ln A^P_{ij}$). All comparative statics derivations are in Online Appendix 3 (and allow the multilateral price term, P_j, to change).[16]

2.3.1. Comparative static 1: extensive margin

Recalling that $\tau_{ij} = t_{ij} + fr_{ij}$, as shown in Online Appendix 3 the model yields that the *ad valorem* tariff-rate elasticity of the extensive margin (EM_{ij}) is given by:

$$\frac{d\ln EM_{ij}}{d\ln t_{ij}} = -\left(\frac{1}{1+\frac{fr_{ij}}{t_{ij}}}\right)\left(\frac{\gamma-(\sigma-1)}{1-\frac{\gamma}{\sigma-1}\eta s_{ij}}\right) < 0 \tag{12}$$

where $s_{ij} = \frac{M_{ij}^{-\eta}}{A^N_{ij}+A^P_{ij}+M_{ij}^{-\eta}}$, that is, the share of endogenous export fixed costs in total export fixed costs.[17] Several key insights are revealed by equation (12). First, as in Chaney (2008), lower tariff rates (t_{ij}) increase the extensive margin; a fall in τ_{ij} directly lowers the export cutoff productivity and increases the number of export firms (M_{ij}). Moreover, as in Krautheim (2012), the increase in the number of exporting firms expands the network effect which further lowers the export cutoff productivity (due to η). Second, while Krautheim's network effect "magnifies" the extensive margin elasticity, it does not make it *endogenous*. However, in our model, the extensive margin elasticity is endogenous to the *level* of exogenous export fixed costs A_{ij}. The lower is either exogenous natural fixed export costs (A^N_{ij}) or policy fixed export costs (A^P_{ij}), the higher is s_{ij}, augmenting the relative importance of the network effect and increasing the (absolute) extensive margin elasticity. Third, the EM elasticity is sensitive to the relative levels of *ad valorem* freight rates and (initial) tariff rates. Hummels and Skiba (2004) found a strong empirical correlation between bilateral distances and measures of fr_{ij}; hence, country-pairs that are closer (and as such have lower fr_{ij}) should have a higher EM elasticity to tariff-rate cuts. Consistent with our introductory quote, the trade-policy elasticity varies with the "particular setting." Finally, note that if the network effect is absent ($\eta = 0$), the extensive margin elasticity is exogenous and simplifies to that in Chaney (2008), $-[\gamma - (\sigma - 1)]$, except for the influence of fr_{ij}/t_{ij}.

2.3.2. Comparative static 2: intensive margin

The *ad valorem* tariff-rate elasticity of the intensive margin (IM_{ij}) is given by:

$$\frac{d\ln IM_{ij}}{d\ln t_{ij}} = -\left(\frac{1}{1+\frac{fr_{ij}}{t_{ij}}}\right)(\sigma-1) < 0 \tag{13}$$

With the exception of the role of fr_{ij}/t_{ij}, this result would identical to that in Chaney (2008) and Krautheim (2012). However, the additive separability in fr_{ij} and t_{ij} implies that the intensive margin elasticity is heterogeneous across country-pairs. A lower *ad valorem* freight-cost factor implies a larger IM elasticity (in absolute terms).

Table 1
Summary of theoretical effects of trade-cost levels on the EM, IM, and trade elasticities.

	Extensive Elasticity	Intensive Elasticity	Trade-Flow Elasticity
Ad Valorem Tariff-Rate Elasticity Effects			
Lower fr_{ij}	+	+	+
Lower A^N_{ij}	+	0	+
Lower A^P_{ij}	+	0	+
Policy Fixed Trade-Cost Elasticity Effects			
Lower fr_{ij}	0	0	0
Lower A^N_{ij}	+	0	+
Lower A^P_{ij}	–	0	–
EIA Coefficient Effects			
Lower fr_{ij}	+	+	+
Lower A^N_{ij}	+	0	+
Lower A^P_{ij}	?	0	?

Notes: See text.

2.3.3. Comparative static 3: aggregate trade flows

As typical to this class of models, the *ad valorem* tariff-rate elasticity of the aggregate trade flow (X_{ij}) is the sum of the previous two elasticities:

$$\frac{d\ln X_{ij}}{d\ln t_{ij}} = -\left(\frac{1}{1+\frac{fr_{ij}}{t_{ij}}}\right)\left[(\sigma-1)+\left(\frac{\gamma-(\sigma-1)}{1-\frac{\gamma}{\sigma-1}\eta s_{ij}}\right)\right] < 0 \tag{14}$$

As common to these types of models, aggregate trade is influenced by *ad valorem* tariff-rate changes via changes in the export cutoff productivity. Hence, the endogeneity of the aggregate trade flow elasticity depends upon the extensive margin elasticity, which as shown above is endogenous to the *levels* of exogenous bilateral policy and non-policy export fixed costs. Moreover, given our *ad valorem* trade-cost function, the intensive margin response to a tariff cut is endogenous to the importance of freight factors relative to initial tariff rates. The first line of the top panel of Table 1 summarizes the qualitative effects of a lower freight factor on the EM, IM, and trade elasticities to tariff-rate changes; lower fr_{ij} implies larger elasticities. The second and third lines of the top panel in Table 1 summarize the qualitative effects of lower natural and policy export-fixed-cost levels, respectively, on the three tariff-rate elasticities just discussed. A lower level of either type of export fixed cost, by causing a rise in s_{ij}, has the same qualitative effect on the three elasticities.[18]

It is useful at this point to note that the non-linear combination of the structural parameters of the model – σ, γ, and η – in the first and third comparative statics precludes identification of any of the individual structural parameters. Moreover, identification of individual structural parameters is compromised by the estimated coefficients being influenced by levels of variable and fixed trade-cost variables. Of all of the tariff-rate elasticities, one might argue that the tariff-rate elasticity of the intensive margin might allow identification of σ, using tariff rate data. However, that elasticity is a function of fr_{ij}, for which data is scarce and of poorer quality than tariff rates.

We close this section noting that – in the absence of endogenous export fixed costs (i.e., $\eta = 0$) – the comparative statics change quantitatively but not qualitatively. The assumed additively separable form for the trade-cost factor, $\tau_{ij} = t_{ij} + fr_{ij}$, is sufficient to generate EM, IM, and trade-flow elasticities endogenous to relative levels of t_{ij} and fr_{ij}.

[16] It will be useful to note that additive separability of components of variable trade costs and of export fixed costs readily gives rise to heterogeneous trade elasticities in levels of trade costs. For instance, for any three variables y, x_1, and x_2, suppose $x = x_1 + x_2$. If the elasticity of y with respect to x is a constant (say, c), then $\partial\ln y/\partial\ln x_1 = c\frac{x_1}{x_1+x_2}$.

[17] As discussed in Online Appendix 3, we assume the share of i's exports in j's total imports is small and the elasticity of j's total tariff revenue to a change in the bilateral tariff rate is small.

[18] Note that s_{ij} takes into account the new equilibrium level of M_{ij} since φ^*_{ij} has changed (except for changes in φ^*_{ij} due to changes in w_j).

2.4. Comparative statics for policy-oriented export fixed costs

The next three comparative statics are related to the effects on the extensive margin, intensive margin, and aggregate trade flow of an exogenous change in bilateral policy fixed export costs.

2.4.1. Comparative static 4: extensive margin

The elasticity of the extensive margin with respect to a one percent change in exogenous bilateral policy export fixed costs (A_{ij}^P) is:

$$\frac{d\ln EM_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\frac{\gamma}{\sigma-1}-1}{1-\frac{\gamma}{\sigma-1}\eta s_{ij}}\right)\left(\frac{A_{ij}^P}{A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\gamma\eta}}\right) < 0. \quad (15)$$

There are two important insights to glean from equation (15). First, equation (15) implies that the lower is the initial *level* of exogenous *non-policy* export fixed costs (A_{ij}^N), the higher (in absolute terms) will be the impact of a one percent change in exogenous policy export fixed costs (A_{ij}^P) on the extensive margin. For example, the impact of an EIA on the extensive margin by lowering A_{ij}^P will likely be higher if the two countries have greater cultural similarities (which likely lower A_{ij}^N). The reason is that a lower level of A_{ij}^N magnifies the elasticity unambiguously by raising both terms in parentheses in the RHS of equation (15). A lower level of A_{ij}^N raises s_{ij}, which magnifies the effect of $d\ln A_{ij}^P$ on the extensive margin as shown in the first parenthetical RHS term. Also, a lower level of A_{ij}^N increases the relative importance of policy fixed export cost changes ($d\ln A_{ij}^P$) captured in the second parenthetical RHS term, further magnifying the elasticity. Moreover, using this result and Comparative Static 1, the effect of an EIA – by lowering both τ_{ij} and A_{ij}^P – on the extensive margin should be unambiguously larger the lower are non-policy export fixed costs A_{ij}^N. These results are summarized in the second line of all three panels of Table 1.

Second, equation (15) suggests a set of different conclusions for initial levels of *policy* export fixed costs (A_{ij}^P). Although a lower level of initial policy export fixed costs, such as common institutional background (common legal origins, etc.), raises s_{ij}, tending to increase the $d\ln A_{ij}^P$ elasticity, a lower level of initial policy export fixed costs lowers the second term in parentheses in equation (15), tending to decrease the $d\ln A_{ij}^P$ elasticity. However, as shown in Online Appendix 3 (section A3.4, Proof), the latter effect dominates as long as we assume, as in Krautheim (2012), that the stability condition $\frac{\gamma\eta}{\sigma-1} < 1$ holds. Hence, as summarized in the third line of the middle panel of Table 1, the policy export fixed cost extensive margin elasticity should decline with lower initial levels of policy export fixed costs. The economic intuition is that a lower A_{ij}^P implies a lower initial level of bilateral policy, or institutional, differences, making the gains from an EIA smaller.

However, as summarized in the third line of the last panel of Table 1, the effect of a lower level of A_{ij}^P on the EIA elasticity is ambiguous theoretically. Although a lower initial A_{ij}^P decreases the policy export fixed cost extensive margin EIA elasticity, a lower initial A_{ij}^P increases the variable-trade-cost extensive margin elasticity (as discussed earlier).

Finally, if we assume no natural export fixed costs ($A_{ij}^N = 0$) and no network externality ($\eta = 0$) as in Chaney (2008), equation (15) simplifies to:

$$\frac{d\ln EM_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\gamma}{\sigma-1}-1\right) < 0 \quad (16)$$

which is exactly the same result as in Chaney (2008).

2.4.2. Comparative static 5: intensive margin

The policy export fixed cost intensive margin elasticity is:

$$\frac{d\ln IM_{ij}}{d\ln A_{ij}^P} = 0 \quad (17)$$

This is analogous to that in Chaney (2008) and is unsurprising. This is summarized in the middle panel of Table 1.

2.4.3. Comparative static 6: aggregate trade flows

Noting the previous two comparative statics, the policy export fixed cost trade-flow elasticity is the same as the policy export fixed cost extensive margin elasticity:

$$\frac{d\ln X_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\frac{\gamma}{\sigma-1}-1}{1-\frac{\gamma}{\sigma-1}\eta s_{ij}}\right)\left(\frac{A_{ij}^P}{A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\gamma\eta}}\right) < 0 \quad (18)$$

The last row of the bottom panel in Table 1 summarizes the ambiguous effects on the EIA extensive margin and trade-flow elasticities of a lower initial level of A_{ij}^P. The tension arises by contrasting the last rows of the top and middle panels. A lower initial A_{ij}^P raises (in absolute terms) the variable-trade-cost elasticities but lowers the export-fixed-cost elasticities. Since an EIA lowers both fixed and variable trade costs, the ambiguity surfaces.

We close this section noting that – in the absence of endogenous export fixed costs (i.e., $\eta = 0$) – the comparative statics change quantitatively but not qualitatively. The assumed additively separable form for exogenous export fixed costs, $A_{ij} = A_{ij}^N + A_{ij}^P$, is sufficient to generate EM and trade-flow elasticities endogenous to relative levels of A_{ij}^N and A_{ij}^P.

3. Econometric model and data sources

In the first section, we discuss the econometric approach. In the second section, we discuss the relationships between our variable natural trade cost, non-policy fixed export cost, and policy fixed export cost theoretical variables and observable proxies suggested in HMR and used for the baseline specifications. In the third section, we discuss the data for the EIA dummies, nominal trade flows, and extensive and intensive margins. In the fourth section, we present the baseline regression specifications.

3.1. Econometric approach

Many of the trade-policy liberalizations in the past 50 years have been bilateral (and plurilateral) EIAs, such as free trade agreements. However, typically EIAs are broad agreements reaching beyond elimination of *ad valorem* tariff rates (which are variable trade costs). They have also lowered policy fixed export costs.[19] For instance, see Horn et al. (2010) on the numerous non-tariff-rate provisions covered in an anatomy of European Union and United States' preferential trade agreements. Thus, EIA liberalizations likely lower t_{ij} (and hence τ_{ijt}) and A_{ijt}^P. Moreover, as noted in Anderson and van Wincoop (2004), empirical *ad valorem* measures of bilateral tariff rates are subject to measurement error. *Ad valorem*-equivalent measures of nontariff barriers and other fixed export costs are worse.[20]

Consequently, many researchers using gravity equations have turned instead to panel data methodologies with dummy variables and fixed effects to find consistent and precise empirical estimates of

[19] This is why earlier we distinguished bilateral fixed export costs associated with "policy," denoted A_{ijt}^P, from bilateral fixed export costs associated with "non-policy," or "natural," factors, denoted A_{ijt}^N.

[20] As noted in Anderson and van Wincoop (2004), "An important theme is the many difficulties faced in obtaining accurate measures of trade costs. Particularly egregious is the paucity of good data on policy barriers. Transport-cost data is limited in part by its private nature. Many other trade costs (that can potentially be reduced by an EIA), such as those associated with information barriers and contract enforcement, *cannot be directly measured at all*" (p. 693; italics added).

the "average treatment effects" of EIAs on trade flows, cf., Baier and Bergstrand (2007), or BB, Anderson and Yotov (2011), Eicher et al. (2012), and Head and Mayer (2014).[21] For instance, BB showed that consistent and precise estimates of partial (treatment) effects of EIAs on bilateral trade flows could be captured using the gravity-equation specification below using ordinary least squares (OLS):

$$\ln X_{ijt} = \alpha + \Theta_{it} + \Psi_{jt} + \psi_{ij} + \beta EIA_{ijt} + v_{ijt} \tag{19}$$

where Θ_{it} is an exporter-year fixed effect, Ψ_{jt} is an importer-year fixed effect, ψ_{ij} is a pair fixed effect, and v_{ijt} is an error term.[22] Equation (19) is commonly referred to as a "fixed effects" model. A key insight of BB was to show methodologically and empirically the importance of the country-pair fixed effect for controlling for the endogeneity of the EIA variable, alongside fixed effects Θ_{it} and Ψ_{jt} to account for exporters' and importers' time-varying GDPs and multilateral price terms.

There are limitations to specification (19). One limitation is that it imposes a common estimated average partial effect (β) for all EIAs. Naturally, EIAs differ in terms of the degree of trade liberalization, with "deeper" agreements expected to have had greater trade liberalization. Historically, several studies have attempted to allow for (*ex post*) heterogeneous EIA effects by introducing instead a multitude of dummies – one for each agreement. However, this approach often leads to weak estimates. The reason is that – unless the EIA is plurilateral with numerous common memberships – there is insufficient variation in the RHS dummy variables. This was the dilemma Tinbergen (1962) faced, leading to the trivial EIA effects of the British Commonwealth and BENELUX economic union.[23] Baier et al. (2014), or BBF, accounted for this – but avoided weak estimates associated with a multitude of dummies – by running a specification including separate dummies for one-way PTAs (OWPTA), two-way PTAs (TWPTA), FTAs, and a dummy combining customs unions, common markets, and economic unions (CUCMECU), due to the limited number of these more integrated EIAs in their sample ending in 2000.[24] Hence, BBF ran the fixed effects model:

$$\begin{aligned}\ln X_{ijt} = {} & \alpha + \Theta_{it} + \Psi_{jt} + \psi_{ij} + \beta_1 OWPTA_{ijt} + \beta_2 TWPTA_{ijt} \\ & + \beta_3 FTA_{ijt} + \beta_4 CUCMECU_{ijt} + v_{ijt}\end{aligned} \tag{20}$$

using OLS. Among other findings, BBF found that deeper economic integration agreements had, as expected, larger average partial effects on bilateral trade flows.

A second limitation of specification (19) (or (20)) is that – even for a given degree of liberalization – the effects of EIAs on trade flows are likely to be heterogeneous across country pairs. In specifications such as equation (19) or (20), this heterogeneity in EIAs' partial effects is captured in the error term, v_{ijt}, which is *assumed* to be uncorrelated with the other right-hand-side (RHS) variables. Yet, the partial effect on trade of EIAs with a given degree of trade liberalization may be heterogeneous due to variable and/or fixed bilateral export costs discussed in section 2. For tractability, suppose EIA_{ijt} represents EIAs with a given degree of trade liberalization. Following Cameron and Trivedi (2005) (p. 774), we can consider the specification:

$$\ln X_{ijt} = \alpha + \Theta_{it} + \Psi_{jt} + \psi_{ij} + \beta_{ij} EIA_{ijt} + v_{ijt} \tag{21}$$

where the partial effect of an EIA on ln X_{ijt} is allowed to be *pair-specific*. The purpose of the comparative statics provided above in sections 2.3 and 2.4 was to provide a theoretical rationale that the effects on trade, intensive margins, and extensive margins from changes in tariff rates and policy-based fixed trade costs associated with formation or dissolution of an EIA (ΔEIA_{ijt}) are sensitive to the *levels* of variable and fixed trade-cost variables. Econometrically, this implies there exists a set of variables Z_{ij} such that:

$$E(\ln X_{ijt} \mid \alpha, \Theta_{it}, \Psi_{jt}, \psi_{ij}, \beta_{ij}, EIA_{ijt}, Z_{ij}) = \alpha + \Theta_{it} + \Psi_{jt} + \psi_{ij} + \beta_{ij} EIA_{ijt} \tag{22}$$

Without knowing the true values of the $\boldsymbol{\beta_{ij}}$, we take expectations over all variables to obtain:

$$\begin{aligned}E(\ln X_{ijt} \mid \alpha, \Theta_{it}, \Psi_{jt}, \psi_{ij}, EIA_{ijt}, Z_{ij}) = {} & \alpha + \Theta_{it} + \Psi_{jt} + \psi_{ij} \\ & + E(\beta_{ij} \mid \alpha, \Theta_{it}, \Psi_{jt}, \psi_{ij}, EIA_{ijt}, Z_{ij}) EIA_{ijt}\end{aligned}$$

We assume that the expected effect of an EIA between i and j, conditioning on all other variables, is given by:

$$E(\beta_{ij} \mid \Theta_{it}, \Psi_{jt}, \psi_{ij}, Z_{ij}) = \beta + b_Z(Z_{ij} - \mu_Z)$$

where $Z_{ij} - \mu_Z$ denotes the *de-meaned* values of Z_{ij}. Suppose, for example, that a single variable, the log of bilateral distance, ln $distance_{ij}$, determines Z_{ij}. Let ln $DIST_{ij}$ = ln $distance_{ij} - \mu_{distance}$. Absent knowledge of β_{ij}, following Cameron and Trivedi (2005) we should estimate instead:

$$\begin{aligned}E(\ln X_{ijt} \mid \alpha, \Theta_{it}, \Psi_{jt}, \psi_{ij}, EIA_{ijt}, Z_{ij}) = {} & \alpha + \Theta_{it} + \Psi_{jt} + \psi_{ij} + \beta EIA_{ijt} \\ & + b(EIA_{ijt} * \ln DIST_{ij})\end{aligned} \tag{23}$$

The main goal of the empirical section of the paper is to identify the variables in Z_{ij}. While incorporating theory-motivated interaction terms is a distinguishing feature of the empirical work in this paper, we will also acknowledge in numerous sensitivity analyses several remaining shortcomings in specification (23). The robustness analyses later will include alternative nontraded goods cutoffs, interactions by type of EIA (e.g., free trade agreement, custom union, etc.), lagged values of *EIA* and interaction terms, accounting for additional controls not explicitly in the theory (especially aspects of development), accounting additionally for tariff rates, accounting for zeros in aggregate trade, and decomposing the sample by sectors.

3.2. Observable proxies for variable and fixed export costs

So what *observable* variables might proxy for the unobservable exogenous variable natural trade costs (fr_{ij}), exogenous non-policy fixed export costs (A^N_{ij}), and exogenous policy fixed export costs (A^P_{ij}) discussed in section 2? Beginning with Tinbergen (1962), the empirical gravity equation literature provides more than 50 years of econometric examination of observable bilateral variables that likely affect trade flows via bilateral trade costs. Typical variables that have surfaced over decades are bilateral distance, measures of religious similarities, and dummy variables for common land border, primary language, legal origin, and colonial history, cf., HMR and Head and Mayer (2014). Up until 2003, this literature has interpreted the channel of influence of these variables on trade flows as the intensive margin. However, three pertinent considerations suggest that some *or all* of these six – what we will term "standard gravity covariates" – might influence *fixed* export costs. First, the trade literature since 2000 has called considerable attention to the theoretical importance of fixed export costs for explaining zeros in trade. Second, Nunn and Trefler (2014) note considerable empirical evidence on the importance of institutions and cultural similarities for explaining international trade, and note that such factors may have a considerable effect on extensive margins of trade. Third, HMR and Egger et al. (2011) (or ELSW) have shown empirically that

[21] Nevertheless, we will discuss the results of a robustness analysis also including *ad valorem* tariff rates, but – due to data constraints – for a shorter panel.

[22] For now, we ignore zero trade flows, allowing a log-linear gravity equation. We address below the robustness analysis that we will provide to account for zeros. See BB and Baier et al. (2014) for theoretical gravity-equation motivation for equation (19).

[23] There were only three countries in each agreement in his sample and only six "1's" in each of the dummy variables.

[24] In this paper, we have extended that data set to 2010, enlarging substantially the number of EIAs with customs unions (CUs), common markets (CMs), and economic unions (ECUs), and so will treat each of those types separately.

Table 2
Expected EIA variables' coefficient signs.

1 Variables	2 Expected Sign Extensive	3 Expected Sign Intensive	4 Expected Sign Total
EIA	+	+	+
EIA*lnDist (fr, A^N)	–	–	–
EIA*ADJ (fr, A^N)	–	+	?
EIA*LANG (A^N)	+	0	+
EIA*RELIG (A^N)	+	0	+
EIA*LEGAL (A^P)	?	0	?
EIA*COLONY (A^P)	?	0	?

Notes: See text.

some of these six variables actually explain the extensive, as well as intensive, margin of trade. However, they also reveal that there are quantitative as well as *qualitative* differences in the impacts of these variables on the two margins of trade. For instance, bilateral distance negatively influences both the probability and volume of trade in both studies. However, contiguity of nations (i.e., sharing a common land border) influences positively the intensive margin, but *negatively* the extensive margin, in HMR and ELSW. Hence, we look to observable standard gravity covariates to explain empirically bilateral variability of fr_{ij}, A_{ij}^N, and A_{ij}^P, key factors in explaining heterogenous EIA effects in the context of our theoretical model.

HMR's Appendix 1 discusses the construction of a number of observable bilateral variables which they classified as *geographic* (including bilateral distances and a dummy for common international land border, termed here adjacency), *cultural* (religious similarity and a dummy for common language), and *institutional* (including dummies for common legal origin and common colonial history). We employ this same categorization.

Following a long-standing tradition, we proxy our *ad valorem* bilateral natural (non-policy) variable trade costs fr_{ij} by distance (ln *DIST*) and adjacency (*ADJ*). Empirical support for distance as a proxy for fr_{ij} is provided in Hummels and Skiba (2004). The adjacency dummy has a rich usage in gravity-equation studies and is commonly interpreted as another factor influencing fr_{ij}. However, while adjacency is likely to lower freight costs and increase the intensive margin of trade, evidence from HMR and ELSW noted above suggests that having a common international land border may create a higher level of natural fixed export costs (A_{ij}^N), i.e., a "border effect." Consequently, in Table 2 we conjecture a negative sign for the EIA interaction term with adjacency for the extensive margin, based upon these previous findings of a negative relationship between adjacency and the probability of a country pair trading.

We associate HMR's bilateral cultural variables – religious similarity (*RELIG*) and common language (*LANG*) – with bilateral *non-policy*, or natural, fixed export costs (A_{ij}^N). As noted in Alesina and Giuliano (2015), most empirical papers adopt the Guiso et al. (2006) definition of cultural variables as "beliefs and values that ethnic, religious, and social groups transmit fairly unchanged from generation to generation." As summarized in Table 2, we expect religious similarity and common language to influence natural fixed export costs between *i* and *j*, but not variable trade costs. Support for this arises from the HMR and ELSW findings that religious similarity and common language had economically and statistically significant effects on the probability of positive trade, but had little or no impact on the level of trade (conditioned on positive trade). We expect religious similarity and common language to increase the extensive margin EIA elasticity and trade-flow EIA elasticity, but have no impact on the intensive margin EIA elasticity based on comparative statics in section 2.

We associate HMR's bilateral institutional variables – common legal origins (*LEGAL*) and common colonial histories (*COLONY*) – with bilateral *policy* fixed export costs (A_{ij}^P). As noted in Alesina and Giuliano (2015), North (1990) defined institutions as "humanly devised constraints that structure human interactions. They are made up of formal constraints (rules, laws, constitutions) …. " As summarized in Table 2, we expect common legal origins and common colonial histories to influence policy fixed export costs, but not variable trade costs. However, in contrast to the cultural proxies, note that the institutional variables' expected effects on the extensive margin and trade-flow EIA elasticities are *ambiguous*. This is because – although a lower level of A_{ij}^P raises s_{ij} tending to increase the variable-trade-cost and fixed-export-cost elasticities – a lower level of A_{ij}^P lowers the relative importance of policy vs. non-policy fixed export costs, diminishing the policy fixed export costs EIA elasticity. The economic intuition is straightforward: if two countries already have a common legal origin or a common colonial history, the gains from an EIA to reduce policy fixed export costs are diminished. Finally, the data for all these bilateral variables are from CEPII.[25]

3.3. Other data for baseline specifications

Other data used are dummy variables for various levels of economic integration agreements (EIAs), nominal aggregate trade flows, intensive margins, and extensive margins.

While several earlier gravity-equation analyses have used dummy variables indicating the presence or absence of an EIA between country pairs for numerous years, there are few publicly available systematic data sets that have multichotomous indexes of EIAs for a large number of country pairs and number of years (i.e., a panel). We use the data set constructed by Scott Baier and Jeffrey Bergstrand and provided at Jeffrey Bergstrand's website, www.nd.edu/(tilde)jbergstr/. The index is defined as: no EIA (0), one-way preferential trade agreement, or *OWPTA* (1), two-way preferential trade agreement, or *TWPTA* (2), free trade agreement, or *FTA* (3), customs union, or *CU* (4), common market, or *CM* (5), and economic union, or *ECU* (6). The definitions are conventional, based upon Frankel (1997), and are defined explicitly in the data set.[26] In this paper, we use 183 countries; Online Appendix 4 lists the EIAs in our sample and (at its end) the countries included. Table 3 provides a decomposition of the data set into types of agreements. Note that the vast majority of observations have no economic integration agreement and less than 6 percent of the observations have FTAs, CUs, CMs, or ECUs. As will be discussed below, initially we use one dummy variable, EIA_{ij}, which includes all FTAs, CUs, CMs, and ECUs. In a robustness analysis, we will include dummy variables for all six types separately.

Nominal disaggregate trade flows are from the United Nations' COMTRADE database for the years 1965, 1970, 1975, 1980, 1985, 1990, 1995, 2000, 2005 and 2010 (using WITS, the World Integrated Trade Solution). The rationale for using only five-year intervals is the same as in BB and BBF, and is explained comprehensively there.[27] As we will also examine EIA effects on the extensive and intensive margins, we need a methodology for a data set with a large number of years and a large number of countries to extract extensive and intensive margins. Fortunately, as used in BBF, Hummels and Klenow (2005), or HK, was the first paper to highlight a tractable method for decomposing transparently the extensive and intensive *goods* margins of trade for a large set of countries' bilateral trade flows using publicly avail-

[25] See http://www.cepii.fr/cepii/en/bddmodele/presentation.asp?id=6 and http://www.cepii.fr/cepii/en/bddmodele/presentation.asp?id=8.
[26] There are several versions of the data set; the one used for this paper is a (2014) extended-to-2011 version of the May 2013 data set.
[27] Due to space constraints here, see BBF, p. 342 and the BBF Online Appendix.

Table 3
Data description.

Integration Index	Count	Percent of Total	Percent of subtotal
0 (None)	567,531	34.8	78.1
1 (1-way PTA)	94,789	5.8	13.0
2 (2-way PTA)	23,184	1.4	3.2
3 (FTA)	25,570	1.6	3.5
4 (Customs Union)	7259	0.4	1.0
5 (Common Market)	5516	0.3	0.8
6 (Economic Union)	2619	0.2	0.4
Subtotal	726,468	–	100.0
Missing observations	905,526	55.5	
Total	1,631,994	100.0	

Notes: Total observations are based upon 183 countries ($183 \times 182 = 33,306$) for 49 years (1962–2010). Missing observations include country pairs with zero trade value and/orone country (or both) of a bilateral pair did not officially exist. See data source at www.nd.edu/~jbergstr.

able disaggregate trade data.[28] Let X_{ijt} denote the value of country i's exports to country j in year t. Following HK, the extensive margin of goods exported from i to j in any year t is defined as:

$$EM_{ijt} = \frac{\sum_{m \in M_{ijt}} X^m_{Wjt}}{\sum_{m \in M_{Wjt}} X^m_{Wjt}} \tag{24}$$

where X^m_{Wjt} is the value of country j's imports from the world in product m in year t, M_{Wjt} is the set of all products exported by the world to j in year t, and M_{ijt} is the subset of all products exported from i to j in year t. Hence, EM_{ijt} is a measure of the fraction of all products that are exported from i to j in year t, where each product is weighted by the importance of that product in world exports to j in year t. Consequently, the HK definition of the extensive margin of trade from i to j corresponds precisely to that used for the comparative statics in section 2.

HK define the intensive margin of goods exported from i to j as:

$$IM_{ijt} = \frac{\sum_{m \in M_{ijt}} X^m_{ijt}}{\sum_{m \in M_{ijt}} X^m_{Wjt}} \tag{25}$$

where X^m_{ijt} is the value of exports from i to j in product m in year t. Thus, IM_{ijt} represents the market share of country i in country j's imports from the world within the set of products that i exports to j in year t. Consequently, the HK definition of of the intensive margin of trade from i to j corresponds precisely to that used for the comparative statics used in section 2.

One of the notable properties of the HK decomposition methodology is that the product of the two margins equals the ratio of exports from i to j relative to country j total imports:

$$EM_{ijt} IM_{ijt} = \frac{\sum_{m \in M_{ijt}} X_{ijmt}}{\sum_{m \in M_{Wjt}} X_{Wjmt}} = X_{ijt}/X_{jt} \tag{26}$$

where X_{jt} denotes j's imports from the world. Taking the natural logs of equation (26) and some algebra yields:

$$\ln X_{ijt} = \ln EM_{ijt} + \ln IM_{ijt} + \ln X_{jt}. \tag{27}$$

[28] Studies have also used country-specific data on individual plants (or firms) to study extensive and intensive *firm* margins of trade liberalization, but such studies have been confined to particular countries because such data is widely known to be much more costly to access and such data sets have not been concorded for international comparisons, as noted in HMR. See Eaton et al. (2011) for a study of French firms, Trefler (2004) for a study of Canada and the United States, and Pavcnik (2002) for a study of Chilean firms. Another relevant theoretical and empirical piece with similar overtones is Arkolakis et al. (2008).

Consequently, the HK decomposition methodology yields that the log of the value of the trade flow from i to j in any year t can be decomposed *linearly* into (logs of) an extensive margin, an intensive margin, and the value of j's imports from the world.[29]

3.4. Baseline specifications

Given all of the above, in the next section we will first estimate:

$$\begin{aligned} \ln EM_{ijt} = {} & \alpha_0 + \Xi_{it} + \Omega_{jt} + \vartheta_{ij} + \alpha_1 EIA_{ijt} + \alpha_2 (EIA_{ijt} * \ln DIST_{ij}) \\ & + \alpha_3 (EIA_{ijt} * ADJ_{ij}) + \alpha_4 (EIA_{ijt} * LANG_{ij}) \\ & + \alpha_5 (EIA_{ijt} * RELIG_{ij}) + \alpha_6 (EIA_{ijt} * LEGAL_{ij}) \\ & + \alpha_7 (EIA_{ijt} * COLONY_{ij}) + \zeta_{ijt} \end{aligned} \tag{28}$$

$$\begin{aligned} \ln IM_{ijt} = {} & \phi_0 + \Pi_{it} + \Lambda_{jt} + \kappa_{ij} + \phi_1 EIA_{ijt} + \phi_2 (EIA_{ijt} * \ln DIST_{ij}) \\ & + \phi_3 (EIA_{ijt} * ADJ_{ij}) + \phi_4 (EIA_{ijt} * LANG_{ij}) \\ & + \phi_5 (EIA_{ijt} * RELIG_{ij}) + \phi_6 (EIA_{ijt} * LEGAL_{ij}) \\ & + \phi_7 (EIA_{ijt} * COLONY_{ij}) + \nu_{ijt} \end{aligned} \tag{29}$$

$$\begin{aligned} \ln X_{ijt} = {} & \beta_0 + \Theta_{it} + \Psi_{jt} + \psi_{ij} + \beta_1 EIA_{ijt} + \beta_2 (EIA_{ijt} * \ln DIST_{ij}) \\ & + \beta_3 (EIA_{ijt} * ADJ_{ij}) + \beta_4 (EIA_{ijt} * LANG_{ij}) \\ & + \beta_5 (EIA_{ijt} * RELIG_{ij}) + \beta_6 (EIA_{ijt} * LEGAL_{ij}) \\ & + \beta_7 (EIA_{ijt} * COLONY_{ij}) + \upsilon_{ijt} \end{aligned} \tag{30}$$

where $\ln DIST_{ij}$ is the (de-meaned) natural logarithm of bilateral distance between i and j, ADJ_{ij} is a dummy assuming the value 1 if i and j share a common international land border (are adjacent) and 0 otherwise, $LANG_{ij}$ is a dummy assuming the value 1 if i and j share a common official language and 0 otherwise, $RELIG_{ij}$ is a measure of religious similarity between countries i and j, $LEGAL_{ij}$ is a dummy assuming the value 1 if i and j share common legal origins and 0 otherwise, $COLONY_{ij}$ is a dummy assuming the value 1 if i and j share a common colonial history and 0 otherwise, and $\Xi_{it}, \Omega_{jt}, \vartheta_{ij}, \Pi_{it}, \Lambda_{jt}, \kappa_{ij}, \Theta_{it}, \Psi_{jt}$, and ψ_{ij} are fixed effects.[30] Because OLS is a linear operator, it follows that $\alpha_0 + \phi_0 = \beta_0, \alpha_1 + \phi_1 = \beta_1, \alpha_2 + \phi_2 = \beta_2$, etc. We would not be able to ensure these relationships if each specification was estimated using a nonlinear operator, such as Poisson Pseudo-Maximum Likelihood (PPML). Table 2 summarizes the expected coefficient signs for each of these specifications.

4. Empirical results

4.1. Baseline empirical results with EIA interactions

Table 4 provides the results of estimating equations (28)–(30) using OLS and panel data for every five years from 1965 to 2010. Columns (2), (4) and (6) provide the expected coefficient signs for the variables' coefficients for the extensive margin, intensive margin, and trade flow equations, respectively, as summarized in Table 2. Coefficient estimates' t-statistics are reported in parentheses. Our EIA variable includes FTAs, customs unions, common markets, and economic unions.

[29] The term $\ln X_{jt}$ will be subsumed in an importer-time fixed effect. The trade data are 5-digit SITC. This is the most disaggregated publicly available data set for bilateral trade flows for a larger number of years and a large number of country pairs, constructed on a consistent basis, necessary for the analysis at hand.

[30] All these variables in the interactions are de-meaned as well.

Table 4
Coefficient estimates for baseline specification.

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
EIA_t	+	0.182***	+	0.104***	+	0.286***
		(4.70)		(3.35)		(7.97)
EIA_t * ln DIST	–	−0.142***	–	−0.087***	–	−0.229***
		(-4.21)		(-3.41)		(-7.30)
EIA_t * ADJ	–	−0.206**	+	0.240***	?	0.034
		(-2.14)		(3.70)		(0.39)
EIA_t * LANG	+	0.174**	0	0.026	+	0.201***
		(2.32)		(0.48)		(2.96)
EIA_t * RELIG	+	0.161**	0	0.085	+	0.245***
		(2.16)		(1.57)		(3.65)
EIA_t * LEGAL	?	−0.139**	0	0.028	?	−0.111**
		(-2.32)		(0.63)		(-2.05)
EIA_t * COLONY	?	−0.362***	0	0.157*	?	−0.205*
		(-2.58)		(1.77)		(-1.68)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.811		0.808		0.906
N		66,940		66,940		66,940

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively. Cutoff for nontraded goods is $1,000,000; this affects the sample size. t-statistics are in parentheses.

First and foremost, Tables 2 and 4 both show that there are 16 coefficient signs for the specifications that have definitive predictions. We note in Table 4 that *15 of these 16 coefficient estimates* with definitive predictions have the expected coefficient sign.

Second, note that the coefficient estimates for *EIA* are positive in the first line in all three columns of Table 4. Hence, at the means of all the bilateral variables, EIAs have significant positive effects on the extensive margin, intensive margin, and aggregate trade flows.

Third, consider the results for the EIA interaction variables' coefficient estimates for the extensive margin. Distance and a common land border have negative effects on the (absolute value of the) extensive margin EIA elasticity, as expected. Cultural variables common language and religious similarity have positive effects on the (absolute) extensive margin elasticity, as expected. These results are consistent with the hypotheses that less distance, not sharing a "border," presence of a common language, and religious similarity decrease the level of natural export fixed costs (A_{ij}^N), increasing the extensive margin EIA elasticity. Moreover, the results suggest that sharing a common legal origin and colonial history tend to reduce the level of policy export fixed costs (A_{ij}^P), lowering the extensive margin EIA elasticity. The latter results suggest that the reduction in export fixed costs from an EIA may be more important than the effects of lower tariffs.

Fourth, consider the results for the intensive margin. Greater distance and not having a common land border likely raise freight costs (fr_{ij}), tending to lower the intensive margin EIA elasticity, as expected. Moreover, except for one interaction variable ($EIA * COLONY$), our proxies for levels of natural and policy fixed export costs (A_{ij}^N and A_{ij}^P, respectively) have no statistically significant impacts on the intensive margin EIA elasticities, as expected.[31]

Fifth, consider the results for trade flows. The coefficient estimates for trade flows are *fully consistent* with those for the two margins and are as expected.

It is useful to show in Table 5 a decomposition of the average treatments on various "treated" groups (ATTs) – that is, a decomposition of the ATTs for various groups – by the various components of the estimated β_{ij}. Although we can do such a decomposition for any pair *ij*, it is more informative to decompose the ATT for "All Countries" with an agreement (as a benchmark) and then decompose the ATT for a group of developing countries with an EIA; in the latter case, we use the South American customs union MERCOSUR. First, recall from section 3.1 that the ATT for a pair of countries *ij* is β_{ij}. The expected value of β_{ij} is composed of an average treatment effect (ATE), which is β in section 3.1, and the ATT adjusts this ATE by a vector of interaction terms, $\beta_Z(Z - \mu_Z)$. For instance, in the special case of only one interaction variable, ln $DIST_{ij}$, equation (23) shows that the ATT would be $\beta + b^*\ln DIST_{ij}$. The full set of interactions for the extensive margin, intensive margin, and overall trade are shown in section 3.4's equations (28)–(30). Second, it is informative to show for each of the two treatment groups, All Countries and MERCOSUR, the relative shares that each interaction term contributes to the ATT. Table 5 reports these shares (that sum to 1) for each of the extensive-margin effect, intensive-margin effect, and trade effect for each of the two treatment groups. We note several insights. For All Countries, the ATE and distance explain most of the ATT. The contributions by Adjacency have the expected signs (opposite for extensive and intensive margins, as expected). Language and Religion have the expected positive signs. Legal and Colony have the expected negative signs (or zero). Importantly, the ATT for trade of 0.67 for MERCOSUR – a customs union for several developing nations – is notably higher than that for the benchmark, 0.55; we discuss this more later. Moreover, for MERCOSUR's trade effect, the ATE and distance play a lesser role (relative to the benchmark), and the variables representing policy- and nonpolicy-export-fixed costs make substantive contributions, consistent with our theoretical hypotheses.

Overall, the results strongly support the model's predictions and comparative statics. Moreover, Fig. 1 illustrates the vast heterogeneity in EIA effects by extensive margin, intensive margin, and aggregate trade flows, implying that *levels* of bilateral trade costs have substantial effects quantitatively on partial EIA effects. Note also that the heterogeneity is more pronounced for *extensive* margin effects, which dominate trade effects, as suggested by our Melitz model.

In the subsequent sections, we pursue sensitivity analyses. Our robustness analyses address several issues for which the results may be

[31] $EIA * COLONY$ was the only variable that unexpectedly had a significant positive effect on the intensive margin elasticity. However, we note that this coefficient estimate was only significant at the 10 percent level.

Table 5
Decomposition of ATTs by contributions.

All Countries				MERCOSUR		
(1) ATT Components	(2) Extensive Shares	(3) Intensive Shares	(4) Trade Shares	(5) Extensive Shares	(6) Intensive Shares	(7) Trade Shares
ATE	0.55	0.46	0.52	0.67	0.26	0.43
DIST	0.42	0.38	0.40	0.47	0.19	0.30
ADJ	−0.05	0.09	0.01	−0.45	0.36	0.03
LANG	0.06	0.01	0.04	0.27	0.03	0.13
RELIG	0.06	0.05	0.06	0.35	0.13	0.22
LEGAL	−0.03	0.01	−0.02	−0.35	0.05	−0.11
COLONY	−0.01	0	0	0.03	−0.01	0.01
Sum	1.00	1.00	1.00	1.00	1.00	1.00

Notes: ATT values for All Countries (with an agreement) for extensive margin, intensive margin, and trade are 0.33, 0.22, and 0.55, respectively. ATT values for MERCOSUR countries for extensive margin, intensive margin, and rade are 0.27, 0.40, and 0.67, respectively.

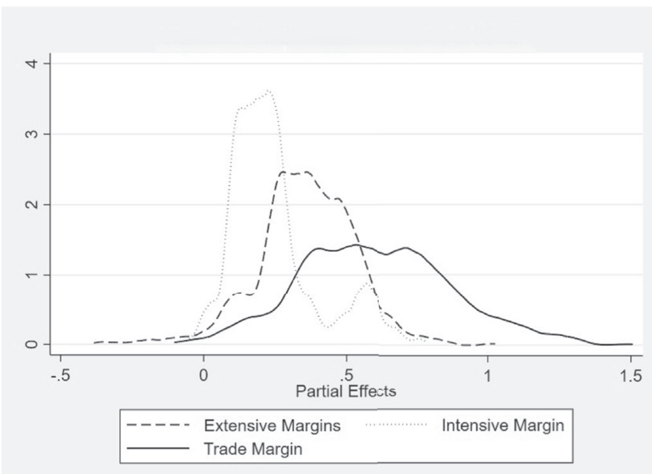

Fig. 1. Heterogeneous EIA effects.

sensitive. First, Kehoe and Ruhl (2013) noted that the effects of EIAs on extensive and intensive margins are sensitive to the choice of "cutoff" values determining traded from nontraded goods; we address this issue. Second, naturally not all EIAs have the same degree of trade liberalization. To address this, we examine the robustness of the results to account for *differing degrees of trade liberalization* by using separate dummies and interactions for one-way PTAs, two-way PTAs, FTAs, customs unions, common markets, and economic unions. Third, we examine whether our interaction results are sensitive to adding lagged effects. Fourth, we have not controlled for various aspects of *development*; we examine the sensitivity of the baseline results to including several variables related to development characteristics of country-pairs. Fifth, despite the limited availability and poor quality of tariff data, we examine the sensitivity of a sub-sample to also including tariff rates. Sixth, we examine the robustness of our specification for aggregate trade flows to inclusion of zeros using a Poisson Pseudo Maximum Likelihood estimator. Seventh, we examine our specifications at the sectoral level.

4.2. Robustness to various nontraded goods cutoffs

As raised in Kehoe and Ruhl (2013), the effects of EIAs on the extensive and intensive margins are sensitive to the choice of "cutoff" values determining traded from nontraded goods. As noted there, to characterize an extensive margin one needs a definition of a nontraded good. Kehoe and Ruhl (2013) show for many trade liberalizations that – using even an absolute cutoff of 50,000 US dollars – there were no extensive margin impacts of EIAs. Using their "relative cutoff" approach, some country pairs' cutoffs for nontraded goods are several millions of US dollars (USD), cf., Table 7 in Kehoe and Ruhl (2013). We have also estimated the results discussed above using cutoffs of 25,000, 50,000, 100,000, 250,000, and 500,000 US dollars, in addition to the 1 million US dollar cutoff used for Table 4. Table 6, for instance, provides the results using the USD 100,000 cutoff. With regard to the statistically significant coefficient estimates, the results between the two tables are fundamentally the same, with the exception of the religion interaction term. In Table 4, religion has a significant impact on EIA's extensive margin effect; by contrast, in Table 6 religion has instead a significant impact on EIA's intensive margin effect.

4.3. Interactions by type of EIA

BBF found that EIA partial effects were smaller for types of agreements with less trade liberalization, as expected. In this section, we investigate whether the interaction terms have the expected effects *by EIA type*. Thus, we determine here empirically whether there are heterogeneous impacts of EIAs at each level of degree of trade liberalization. Consequently, our specifications for extensive margin, intensive margin, and trade flows are expanded to include dummy variables for all six types of EIAs *and* all their interactions. This results in *126 coefficient estimates* for each cutoff value explored.

Because of this very large number of coefficient estimates and t-statistics, the complete set of results by each EIA type including the interaction terms is presented in Online Appendix 5, Table 1. For brevity, we present here two representative sets of results in Tables 7 and 8. Table 7 provides the results for FTAs extracted from Online Appendix 5, Table 1, using the USD 1 million cutoff; there are 21 coefficient estimates (and t-statistics) presented. In Table 7, all the interaction terms have coefficient estimate signs consistent with expectations (when designated). Greater distance diminishes both the intensive and extensive margin elasticities as in Table 4. Adjacency increases the intensive margin elasticity and decreases the extensive margin elasticity as before (though the latter's coefficient estimate is not significant). Common language and religious similarity have no material effect on the intensive margin elasticities and have significant positive effects on the extensive margin elasticities, as expected. Common legal origins and common colonial history have no material effects on the intensive margin elasticities and have significant negative effects on the extensive margin elasticities, as expected.

Table 8 provides the results for customs unions extracted from Online Appendix 5, Table 1; again, there are 21 coefficient estimates (and t-statistics). Consistent with Table 7, all the interaction terms have coefficient signs consistent with expectations (when designated). The only notable difference is that the coefficient estimates for *CU* ∗ *LEGAL* for aggregate trade and the extensive margin are positive, but they are statistically insignificant. A more detailed review of Table 1 in Online Appendix 5 shows that the results are largely the same for all six EIA types.

Table 6
Coefficient Estimates using Alternative Nontraded Goods Cutoff.

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
EIA_t	+	0.012	+	0.227***	+	0.239***
		(0.36)		(7.44)		(6.89)
EIA_t * ln DIST	–	−0.062**	–	−0.133***	–	−0.195***
		(-2.21)		(-5.40)		(-6.60)
EIA_t * ADJ	–	−0.284***	+	0.210***	?	−0.074
		(-3.51)		(3.26)		(-0.86)
EIA_t * LANG	+	0.150**	0	0.068	+	0.218***
		(2.42)		(1.32)		(3.40)
EIA_t * RELIG	+	0.084	0	0.224***	+	0.308***
		(1.44)		(4.47)		(5.01)
EIA_t * LEGAL	?	−0.084*	0	−0.046	?	−0.130**
		(-1.70)		(-1.14)		(-2.48)
EIA_t * COLONY	?	−0.546***	0	0.224***	?	−0.322***
		(-5.67)		(2.65)		(-2.88)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.810		0.788		0.891
N		99,637		99,637		99,637

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively. Cutoff for nontraded goods is $100,000; this affects the sample size. t-statistics are in parentheses.

Table 7
Coefficient estimates for free trade agreements.

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
FTA_t	+	0.127***	+	0.127***	+	0.254***
		(3.08)		(3.89)		(6.63)
FTA_t * ln DIST	–	−0.163***	–	−0.044*	–	−0.207***
		(-4.66)		(-1.65)		(-6.28)
FTA_t * ADJ	–	−0.184	+	0.304***	?	0.120
		(-1.56)		(3.95)		(1.07)
FTA_t * LANG	+	0.149*	0	0.044	+	0.193**
		(1.71)		(0.72)		(2.45)
FTA_t * RELIG	+	0.264***	0	−0.053	+	0.211***
		(3.19)		(-0.88)		(2.79)
FTA_t * LEGAL	?	−0.165**	0	0.069	?	−0.096
		(-2.50)		(1.45)		(-1.61)
FTA_t * COLONY	?	−0.302**	0	0.132	?	−0.170
		(-2.05)		(1.30)		(-1.41)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.811		0.810		0.906
N		66,940		66,940		66,940

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively. Cutoff for nontraded goods is $1,000,000; this affects the sample size. t-statistics are in parentheses.

We also estimated the specifications above using the alternative cutoff of USD 100,000. The results, analogous to those in Table 1 of Online Appendix 5, are presented in Table 2 of Online Appendix 5. For brevity, we will not provide a detailed discussion of these results as they are quite similar to those using the USD 1 million cutoff. Regarding Online Appendix 5, Table 2, there are few changes relative to the Online Appendix 5, Table 1 results that cannot be explained by the fact that – with a lower nontraded good cutoff – there are larger impacts of the interaction variables on intensive margin EIA effects relative to extensive margin EIA effects.

Finally, in Figs. 2–5 (at the end) of Online Appendix 5, we present density plots of the trade, intensive margin, and extensive margin heterogeneous partial effects separately for FTAs, customs unions, common markets, and economic unions, using the USD 1 million nontraded good cutoff. The distinguishing feature of comparing the results is that the average extensive margin effects are larger than the average intensive margin effects for lower levels of trade liberalization, that is, FTAs and customs unions. For common markets and economic unions, the average intensive margin effects are larger than the average extensive margin effects. The economic explanation for this result is intuitive. Deeper levels of economic integration have already likely overcome export fixed costs in earlier stages of integration. Consequently, it is the less liberalized EIAs – such as FTAs and customs unions – where the benefits of having common cultural and institutional factors influence to a larger extent the effect of an FTA or CU by reducing export fixed costs.

Table 8
Coefficient estimates for customs unions.

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
CU_t	+	0.177	+	0.474***	+	0.651***
		(0.79)		(3.11)		(4.07)
CU_t * ln DIST	–	−0.200	–	0.081	–	−0.119
		(-1.49)		(0.85)		(-1.25)
CU_t * ADJ	–	−0.190	+	0.267**	?	0.077
		(-1.07)		(1.97)		(0.52)
CU_t * LANG	+	0.607***	0	0.043	+	0.650***
		(2.94)		(0.30)		(4.33)
CU_t * RELIG	+	0.241	0	0.048	+	0.289**
		(1.57)		(0.34)		(2.17)
CU_t * LEGAL	?	0.043	0	0.042	?	0.085
		(0.35)		(0.45)		(0.80)
CU_t * COLONY	?	−1.119***	0	−0.024	?	−1.143***
		(-3.26)		(-0.13)		(-4.25)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.811		0.810		0.906
N		66,940		66,940		66,940

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively. Cutoff for nontraded goods is $1,000,000; this affects the sample size. t-statistics are in parentheses.

4.4. Lagged effects

BB and BBF estimated treatment effects also allowing for lags. We augmented the model using equation (30) for aggregate trade flows to include five-year lags of the RHS variables, which reduced the sample size. With the exception of two coefficients, the results are insensitive to including lags. For brevity of space, the results for aggregate trade flows are presented in Table 3 of Online Appendix 5.

4.5. Development-related EIA interactions

While section 3 motivated the relevance of HMR's geographic, cultural and institutional bilateral variables for our variable and export-fixed trade costs, there is a notable absence of control variables for aspects of development. In this section, we address several development-related variables that are bilateral combinations of country-specific development characteristics.[32]

First, we consider a bilateral dummy variable that indicates whether or not both countries are members of the World Trade Organization (WTO), called $WTO - BOTH$; this dummy assumes the value 1 (0) if both countries are members of the WTO (0 otherwise). HMR showed that $WTO - BOTH$ has a positive effect on the level of two countries' trade and the probability both countries traded. Both countries being in the WTO likely lowers trade costs. First, common WTO membership likely reduces tariff rate levels (via most-favored-nation (MFN) rules). Recalling Comparative Static 2, this likely reduces the level of t_{ij}, which would tend to decrease the EIA intensive-margin elasticity. Second, common WTO membership likely reduces policy export-fixed costs. However, as discussed above and summarized in the bottom panel of Table 1, $WTO - BOTH$ then could have a positive or negative effect on the EIA extensive-margin and trade elasticities.

Second, we consider differences in the levels of democratic institutions in the two countries, using the POLITY2 index. Such differences may impose higher policy export fixed costs on the pair. Consequently, the expected signs for the EIA extensive-margin and trade elasticities are ambiguous as discussed above. We define the variable *DPOLITY* as the absolute difference in the levels of the POLITY2 index.[33]

Third, we consider differences between and levels of the two countries' per capita GDPs, since per capita GDP is the most common measure of development status. Greater differences in two countries' per capita GDPs are likely to raise trade costs, especially export fixed costs. However, it is unclear *a priori* whether this will affect natural or policy export fixed costs. This then suggests ambiguous effects *a priori* on the EIA extensive-margin and trade elasticities. Similarly, the expected effects of the levels of exporter and importer per capita GDPs on the extensive-margin and trade elasticities are ambiguous; higher levels of per capita GDPs may reduce policy or non-policy export-fixed costs, implying ambiguous expected EIA effects on the extensive-margin and trade elasticities.[34]

Table 9 provides the results from estimating this expanded set of regressions. Unfortunately, incorporating this larger set of variables including governance indicators reduces the sample size (due to data availability); for our benchmark USD 1,000,000 cutoff, the sample size falls from 66,940 to 58,733. Consequently, we first reproduce the results from the specification in Table 4 using the smaller sample so that one can compare these to the development-variables-expanded regressions.[35]

Several points are worth noting. First and most importantly, regardless of extensive-margin, intensive-margin, or aggregate trade results, the baseline interactions' coefficient estimates (and t-statistics) do not change materially when compared to the respective coefficient estimates of the augmented regressions. Second, the coefficient estimate for *EIA* interacted with $WTO - BOTH$ has the expected sign for the intensive margin and is statistically significant. This coefficient estimate for the intensive-margin EIA effect is negative, consistent with Comparative Static 2; the coefficient estimates for the extensive-margin and trade EIA elasticities could not be signed *a priori*. Third, *DPOLITY* had negative and statistically significant effects on extensive-margin as well as trade

[32] We thank a referee for these suggestions.

[33] The data sources is the Polity IV Project, cf., http://www.systemicpeace.org.

[34] GDP and population data are from CEPII.

[35] Comparable results are available using the USD 100,000 cutoff in Online Appendix 5, Table 4.

Table 9
Coefficient estimates for robustness analysis.

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4)	(5) Expected Sign Intensive	(6) Intensive	(7)	(8) Expected Sign Trade	(9) Trade	(10)
EIA_t	+	0.175***	0.161***	+	0.119***	0.044	+	0.294***	0.206***
		(4.39)	(3.45)		(3.71)	(1.21)		(7.96)	(4.82)
EIA_t * ln DIST	–	−0.156***	−0.144***	–	−0.101***	−0.110***	–	−0.258***	−0.254***
		(-4.50)	(-4.11)		(-3.83)	(-4.14)		(-8.03)	(-7.75)
EIA_t * ADJ	–	−0.188*	−0.212**	+	0.229***	0.267***	?	0.041	0.055
		(-1.92)	(-2.10)		(3.45)	(4.08)		(0.45)	(0.59)
EIA_t * LANG	+	0.262***	0.276***	0	−0.018	0.025	+	0.243***	0.301***
		(3.34)	(3.49)		(-0.32)	(0.43)		(3.48)	(4.19)
EIA_t * RELIG	+	0.096	0.113	0	0.133**	0.091	+	0.229***	0.204***
		(1.23)	(1.43)		(2.35)	(1.56)		(3.25)	(2.83)
EIA_t * LEGAL	?	−0.189***	−0.210***	0	0.035	0.071	?	−0.154***	−0.138**
		(-2.97)	(-3.19)		(0.75)	(1.49)		(-2.71)	(-2.37)
EIA_t * COLONY	?	−0.404**	−0.376**	0	0.157	0.103	?	−0.247*	−0.273*
		(-2.53)	(-2.31)		(1.61)	(1.02)		(-1.81)	(-1.96)
EIA_t * WTO-BOTH$_t$	?		0.465***	–		−0.149***	?		0.315***
			(6.49)			(-2.59)			(4.85)
EIA_t * DPOLITY$_t$	?		−0.017***	?		0.007*	?		−0.010**
			(-3.40)			(1.80)			(-2.14)
EIA_t * Difference in ln PCGDP$_t$	?		0.063**	?		−0.066***	?		−0.003
			(2.06)			(-2.85)			(-0.11)
EIA_t * ln Exporter PCGDP$_t$	?		−0.085***	?		0.116***	?		0.031
			(-3.42)			(6.03)			(1.39)
EIA_t * ln Importer PCGDP$_t$	?		−0.072***	?		0.062***	?		−0.010
			(-2.95)			(3.30)			(-0.45)
Fixed Effects:									
Exporter-Year		Yes	Yes		Yes	Yes		Yes	Yes
Importer-Year		Yes	Yes		Yes	Yes		Yes	Yes
Country-Pair		Yes	Yes		Yes	Yes		Yes	Yes
R^2		0.816	0.816		0.799	0.799		0.904	0.904
N		58,733	58,733		58,733	58,733		58,733	58,733

Notes: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$, respectively. Cutoff for nontraded goods is $1,000,000; this affects the sample size. t-statistics are in parentheses.

EIA elasticities; one surprise is that it had a significant positive effect on the intensive-margin EIA elasticity. Finally, differences between and average levels of country-pairs' per capita GDPs had no statistically significant effect on the aggregate trade EIA elasticity. However, these EIA interactions did affect extensive-margin and intensive-margin EIA elasticities. For instance, the lower is two countries' average per capita incomes, the *higher* is the extensive-margin effect of an EIA (with statistical significance). Yet, the lower is two countries' average per capita income, the *lower* is the intensive-margin effect of an EIA (with statistical significance). We return later in section 5 to explaining variation in the heterogeneous partial EIA effects in terms of levels of development and provide more clarity on their relationship.

4.6. Tariff rates

As discussed earlier, most EIAs reduce tariff rates, but also liberalize all types of non-tariff measures and policy-export-fixed costs that are very difficult to quantify.[36] Thus, dummy variables better capture "treatment." Nevertheless, one may question whether all of the effects of tariff-rate reductions are captured by the EIA dummy. To address this, we provide a robustness analysis to determine the sensitivity of our estimates to the additional inclusion of tariff rates. The most daunting restriction of such an analysis is that data on average bilateral tariff rates for a large number of countries for a large number of years is quite limited. Nevertheless, we use the World Integrated Trade System (WITS) data base to find average tariff levels for as many years and country-pairs as possible. Unfortunately, data is only available on a consistent basis for 1990–2010, eliminating 25 years from our analysis (especially the early years when initial tariff rates were higher). Using our benchmark USD 1,000,000 cutoff, our sample size falls *by 78 percent* from 58,733 to 12,892.[37]

This robustness analysis was conducted in three stages. First, we re-ran the extensive-margin, intensive-margin, and aggregate trade specifications in Table 4, but now for the reduced sample of $N = 12,892$. Second, we re-ran the specification in Table 9 expanded to include EIA-development-variable interactions for the reduced sample, also including an EIA interaction with the (de-meaned log of the) average level of the gross bilateral tariff rate, $EIA_{ijt} * \ln TAR_{ijt}$. Third, we expanded the previous specification to include the (log of the) gross bilateral tariff rates, $\ln t_{ijt}$ (where $t_{ijt} > 1$, as earlier), and its interactions with the full complement of covariates being used. Table 5 (Table 6) in Online Appendix 5 provides the main results for the benchmark USD 1,000,000 (alternative USD 100,000) cutoff.

Several points are worth summarizing. First, due to the severe abbreviation of the number of years and the restriction to the post-1989 period, the number of statistically significant coefficient estimates in our baseline model declines dramatically. For instance, for aggregate trade and the extensive margin with the reduced sample and the same specification as in Table 4, only EIA_{ijt}, $EIA_{ijt} * RELIG_{ij}$, and $EIA_{ijt} * COLONY_{ij}$ have statistically significant coefficient estimates, though they do maintain the expected signs. Second, in the specification adding EIA-development-variable interactions, the previously mentioned coefficient estimates are unchanged, but *none* of the EIA-development-variable interactions' nor the new $EIA_{ijt} * \ln TAR_{ijt}$ variable's coefficient

[36] We recommend the reader to the Baier-Bergstrand EIA Data Base at www3.nd.edu/(tilde)jbergstr/. Hyperlinks of any country-pair's cell entry will link the reader to a PDF of its treaty, allowing the reader to see the extent of non-tariff provisions and reductions in export fixed costs associated with that agreement.

[37] Using our USD 100,000 cutoff, the sample falls by 81 percent from 99,637 to 18,720.

estimates are statistically significant. Third, adding ln t_{ijt} and all the associated tariff-rate interactions leaves the previous coefficient estimates materially unchanged. Fourth, of all the new variables added – ln t_{ijt} and all its associated interactions – only ln t_{ijt} is statistically significant for the extensive margin, intensive margin, and aggregate trade; only one tariff-rate interaction's coefficient estimate is statistically significant.[38]

Finally, we discuss the values of the ln t_{ijt} coefficient estimates. One would expect the coefficient estimates on ln t_{ijt} to be negative. In fact, for aggregate trade and the extensive margin, the coefficient estimates are negative, but range between −1.5 and −0.5. However, the tariff-rate elasticity estimate for the intensive margin is *positive*. In the context of our model, theory suggests this value should be negative. We attribute this counter-intuitive result to the weak power of the estimates, due largely to our argument that the tariff data are of insufficient sample size and quality to be useful for this exercise, as well as the presence of the EIA dummies which capture most of the effects of the tariff-rate reductions. Nevertheless, the results are presented in the Online Appendix 5's Tables 5 and 6.[39]

4.7. Zeros

Up to now, we have used only positive trade flow values in our empirical analysis. One reason is this allows us to use OLS for our estimator. A second reason is that, as in Hummels and Klenow (2005), OLS also enables a decomposition of the overall trade flows into the extensive-margin and intensive-margin variables, which are important for our analysis. However, the reader may be curious as to the sensitivity of our empirical results to issues raised in Santos Silva and Tenreyro (2006). One issue raised there is the exclusion of zeros (in trade flows). The other issue is that – due to Jensen's inequality – OLS may be an inferior estimator to Poisson Pseudo-Maximum Likelihood (PPML), which has been adopted widely. We provide a robustness analysis of our results to PPML, noting that – because PPML is not a linear operator – we can only do the analysis for aggregate trade flows and not the decomposition to extensive and intensive margins. Yet, because we conduct the sensitivity analysis only for aggregate trade flows, we are able to use a much larger sample of 152,550 positive flows in the first specification and 232,358 positive and zero flows in the second specification.

Table 7 in Online Appendix 5 provides two sets of coefficient estimates using PPML. We summarize here several points worth noting. First, both specifications yielded coefficient estimate signs largely consistent with expected signs and with previous results. At the means of all the interaction variables for positive flows, the EIA effect is 0.10. The EIA interaction with distance has a negative effect and the EIA interaction with religion has a positive effect (both as expected). The EIA interaction effects with *COLONY* and *LEGAL* are positive – different from earlier – but are statistically insignificant; recall, the expected signs for both of those variables are ambiguous. The only variable with a significant unexpected sign for the coefficient estimate was the EIA interaction variable with common language; it was negative here. Second, and perhaps more important, the coefficient estimates were *virtually identical* between the PPML specifications with just positive trade flows and with positive and zero flows. Thus, as PPML gravity equation coefficient estimates often differ from their OLS counterparts (cf., Bergstrand et al. (2015)), the result that the two PPML specifications had nearly identical results suggested the results were robust to including zeros.

4.8. Sectoral analysis

Our final robustness analysis examines trade flows disaggregated by single-digit Standard International Trade Classification (1-digit SITC). The first robustness analysis in this subsection was to determine if disaggregation by 1-digit SITC provided any further insights into the baseline specification's interaction terms; subsequent robustness analyses extended the model to include development-variable-EIA interactions and then tariff rates and tariff-rate interactions.

As noted above, one of the major limitations of exploring the impact of EIAs including tariff rates is the reduced sample size and the restriction to the post-1989 period. We faced a similar restriction in this robustness analysis, anticipating as before a significant reduction in explanatory power of the interaction variables once tariff rates were included due to the small sample size and sample time period. Because of the large number of sectors (10) and the multiple specifications, for brevity the results are presented in Online Appendix 5, Table 8.[40]

We summarize the main findings of this robustness analysis. First, for the baseline specifications, the average partial effects at the means of the interaction variables vary around the average partial effects using aggregate data. For instance, for trade flows, the partial effect at the means for aggregate trade is 0.145 (see Online Appendix 5, Table 5); for disaggregated sectors, eight significant partial effects range from 0.130 to 0.288 (with two insignificant effects of 0.056 and −0.106). Second, with the reduced sample size due to including tariff data, most of the interaction variables' coefficient estimates are statistically insignificant (as discussed earlier in section 4.6). Third, when interaction variables' coefficient estimates are statistically significant, they generally have the same signs as found in earlier tables. We also ran baseline specifications not restricted by the tariff-rate data limitations. In particular, for manufacturing sectors 5–8 (chemicals, manufactures classified by material, machinery and transport equipment, and miscellaneous manufactures), the results were quite similar to the aggregate trade flow findings.[41]

5. Development status and welfare implications

This section has two goals. First, we provide evidence of the variation in the heterogeneous EIA partial effects across country pairs with agreements and discuss how our framework can provide improved *ex ante* predictions of EIA partial effects. We then show that the variation in EIA partial effects – using the estimates with only the baseline specifications' interaction variables – is significantly related to the level of development (as measured by average per capita GDP); less developed countries *tend* to have higher EIA partial effects. Second, we link the EIA partial effects to estimates of general equilibrium welfare effects. We show that EIA partial effects explain a very large portion of the general equilibrium welfare effects of an EIA.

[38] For aggregate trade, ln $t_{ijt} * RELIG_{ij}$ is statistically significant with the expected negative coefficient.

[39] Were the tariff-rate estimates just discussed of higher quality, one might argue that we would be able to infer estimates of the structural parameters of the model, i.e., σ, γ and η. However, careful examination of Comparative Statics 1–3 reveals that – in the cases of 1 and 3 – the nonlinear combination of the structural parameters precludes identification and – in the case of Comparative Static 2 – the absence of quality data on *fr* precludes identification of σ.

[40] As noted, we estimated the regressions by sector and including tariff rates alongside EIA dummies, EIA interactions, and exporter-year, importer-year, and pair fixed effects. Only in the case where pair fixed effects and EIA dummies and their interactions were excluded were we able to find tariff-rate elasticity estimates in the range of −5 to −17. Once pair fixed effects were included, the tariff-rate elasticity estimates were as described above.

[41] When the data set was not constrained by the tariff data and was much larger (i.e., the disaggregate analogues to the baseline specification), then the sectoral results were much better. For instance, for manufactures sectors SITC 5, 6, 7, and 8 disaggregate trade flows, for every EIA interaction variable's coefficient estimate that was statistically significant, it had the *same sign* as in the aggregate trade flow results in Table 4, with the exception of only one coefficient estimate.

5.1. Heterogeneous EIA (partial) effects and development status

5.1.1. Ex ante analysis

As discussed earlier in the paper, one of the problems with most previous empirical gravity-equation studies of the partial effects of EIAs on trade flows is that there is little variation in the RHS dummy variable so that estimates are weak. A second problem is that they provide only *ex post* estimates for those pairs. Subsequent cross-section analyses using broad samples of country-pairs and EIA dummies provided better "average" effects, but suffered from endogeneity biases. Beginning with Baier and Bergstrand (2007), comprehensive trade-flow and EIA dummy panel data sets allowed more precise and unbiased estimates of EIA partial effects on trade. Moreover, these results suggested a partial effect that could be used *ex ante* for any new agreement. Baier et al. (2014) extended this work by introducing multiple dummy variables for different levels of economic integration and by differentiating partial effects by extensive margin, intensive margin, and (overall) trade.

This study is different by estimating the effects distinguished by proxies for levels of variable and export-fixed trade costs. This generates, as shown in Fig. 1 (and Online Appendix 5's Figs. 2–5), heterogeneous EIA partial effects. In principle, one can predict *ex ante* the partial EIA effect for any pair of countries without an EIA using the average partial effect (for that type of EIA) and time-invariant information on their bilateral distance and statuses of adjacency, religious similarity, language similarity, common legal origin, and common colonial history. Thus, the analysis in this paper provides a theoretically motivated framework for a more precise *ex ante* estimate of the partial effect of a potential EIA than simply the homogenous estimated EIA partial effects implied in Baier and Bergstrand (2007) and Baier et al. (2014).

5.1.2. The role of development status

In the context of the theory and empirical results, we have argued that the heterogeneity in EIA partial effects on trade (and the margins) can be explained by variable and fixed trade costs. Our findings suggest that higher EIA trade effects should be associated with lower natural variable and fixed trade costs but with higher policy fixed export costs. One surprising finding in section 4.5 was that – once such variable and fixed trade cost variables were included – differences between and average levels of per capita GDPs had no material marginal effect on the heterogeneous EIA effects.[42] Does this finding for aggregate trade flows imply that the level of development is immaterial for explaining the aggregate-trade heterogeneous partial EIA effects? Not necessarily. It may well be that developing countries have much higher policy fixed export costs (such as higher government border-crossing costs and weaker port infrastructures), and the effect of the level of development may be influencing the heterogeneous partial EIA effects via this channel. We now explore this possibility.

Using the 2460 estimates of partial EIA effects (for 2460 country pairs) from our earlier analysis, we regressed these EIA estimates on a constant and the logarithm of the average per capita GDPs of each country pair. The regression result was:

$$\widehat{\beta_{ij}} = 0.88 - 0.03\ln(PCGDP_i + PCGDP_j) \tag{31}$$

where the coefficient estimate of 0.03 was statistically significant at the 1 percent level (standard error of 0.005). This result implies that a 10 percent lower average per capita income of a country pair is associated with a 0.30 increase in the EIA partial effect. Since the average partial effect is around 0.50, this suggests that a 10 percent reduction in average per capita GDP can increase the EIA partial effect by *60 percent* (=(0.30/0.50)×100). Based on this evidence, less developed countries are likely to benefit *much more* than developed countries from an EIA, likely due to the substantive reduction of policy export fixed costs.

[42] However, these variables did materially influence extensive- and intensive-margin elasticities.

5.2. Welfare implications

How important *quantitatively* are such heterogeneous partial EIA impacts for overall welfare gains? First, we show in the context of our model the relationship between the general equilibrium welfare impact (labeled $d\ln V_{ij}$), the partial bilateral effect of an EIA, and the bilateral trade share. We provide econometric evidence that $d\ln V_{ij}$ is explained well by these two terms, which may be much easier to measure for a large number of countries and large number of EIAs. Second, we provide a robustness analysis showing that the probability of a country pair having an EIA – which is a proxy for the pair's welfare gain from the EIA, as suggested in Baier and Bergstrand (2004) – is also well explained by the partial impact.

5.2.1. General equilibrium welfare vs. partial effects

We follow the supplementary appendix of Redding (2011) to derive welfare in the context of our model. If we assume exporter network spillovers only apply internationally (not to intranational trade), then welfare in our model is identical to that in Redding (2011) and the welfare effects of trade liberalizations are captured by the same two sufficient statistics discussed in Arkolakis et al. (2012), as shown in Online Appendix 6.

Following Costinot and Rodriguez-Clare (2014), we know by Shephard's Lemma for a small change in trade costs that:

$$d\ln V_j = d\ln w_j - d\ln P_j = d\ln w_j - \ln\sum_{k=1}^{N}\lambda_{kj} d\ln p_{kj}. \tag{32}$$

where λ_{kj} is the share of country j's total expenditures (gross output) spent on goods from country k. In our context, it is useful to rewrite equation (32) as:

$$d\ln V_{ij} = -\lambda_{ij} d\ln p_{ij} + d\ln w_j - \sum_{k\neq i}^{N}\lambda_{kj} d\ln p_{kj} \tag{33}$$

where $d\ln V_{ij}$ denotes the (log) change in country j's welfare from an EIA with country i. Note that equation (33) reveals that welfare changes in j from a bilateral trade-cost change can be decomposed (as conventionally) into a partial effect (first RHS term) and general equilibrium effects (second and third RHS terms). The second RHS term is the welfare-change effect from changes in income and the third RHS term is the welfare-change effect from changes in j's multilateral price term (excluding p_{ij}).

Given an EIA can lower both variable and fixed trade costs, equation (33) can be written as:

$$d\ln V_{ij} = -(1/\gamma)\lambda_{ij}\widehat{\beta_{ij}} + \left(d\ln w_j - \sum_{k\neq i}^{N}\lambda_{kj} d\ln p_{kj}\right) \tag{34}$$

where $\widehat{\beta_{ij}}$ is the trade-flow effect of the bilateral trade-cost change and $(1/\gamma)\widehat{\beta_{ij}} = d\ln p_{ij}$. Defining the general equilibrium effects as $\chi_{ij} \equiv d\ln w_j - \sum_{k\neq i}^{N}\lambda_{kj} d\ln p_{kj}$, we can write:

$$d\ln V_{ij} = -(1/\gamma)\lambda_{ij}\widehat{\beta_{ij}} + \chi_{ij}. \tag{35}$$

Equations (34) and (35) decompose the welfare effect into the partial effect (the first RHS term) and general equilibrium effects (the second RHS term).[43] The intuition is straightforward: $\widehat{\beta_{ij}}$ is the bilateral trade effect of the liberalization, λ_{ij} measures the relative importance of the trading partner, and γ influences the effect on welfare with higher (absolute) γ diminishing the welfare gain.

Our goal in the remainder of this section is to demonstrate empirically that the bulk of variation in $d\ln V_{ij}$ can be explained by variation

[43] In one set of estimates later, unobservable variation can be captured by an import j fixed effect and random error, since by definition χ_{ij} varies only across importers.

in λ_{ij} and $\hat{\beta}_{ij}$, i.e., the partial effect contributors. Although $d\ln V_{ij}$ is not directly observable, it can be estimated. Online Appendix 7 describes in detail how baseline (b) and counterfactual (c) values of V_{ij} are calculated for 1358 bilateral EIA (ij) liberalizations. However, we summarize the process here. First, based upon equations (21) and (30), the baseline bilateral trade cost for any one of these pairs is $\widehat{\psi_{ij}} + \hat{\beta}_{ij}EIA_{ijt}$. Following Baier et al. (2017) and Head and Mayer (2014), we can use our structural gravity equation framework to generate multilateral outward (exporter) price and multilateral inward (importer) price terms from which wage rates w_j and nominal gross outputs Y_j ($= w_jL_j$) can be determined. Following Head and Mayer (2014), we assume $\gamma = 5$. In the baseline scenario, we generate the matrix of trade flows X_{ij}^b (including X_{jj}^b) using the $\widehat{\psi_{ij}} + \hat{\beta}_{ij}EIA_{ijt}$, imputed multilateral outward and inward price terms, and actual nominal gross outputs Y_j. The latter were obtained from the World Input-Output Data (WIOD) base for 2005; data for 61 countries allowed examining 1358 EIA liberalizations. Hence, initial w_j^b were set equal to per capita nominal gross outputs (Y_j^b/L_j). This yielded baseline bilateral trade costs, international and intra-national trade shares, wage rates, nominal incomes, and importer CES price indexes (P_j^b). Consequently, we solved for baseline w_j^b/P_j^b, which captures the initial value of V_{ij}^b.

Computation of the counterfactual welfare level, V_{ij}^c is then straightforward. For each of the 1358 (ij) bilateral liberalizations, we remove the EIA, eliminating the partial (direct) effect of the EIA on X_{ij}, generating a set of counterfactual bilateral trade costs, $\widehat{\psi_{ij}}$. Using the new counterfactual trade costs, we compute the counterfactual multilateral outward price terms, multilateral inward price terms, nominal wage rates, and nominal gross outputs. These variables are then used to generate a set of counterfactual international and intra-national trade flows, which are then used to determine a new set of multilateral outward price terms, multilateral inward price terms, nominal wage rates, and nominal gross outputs. We iterate using a dampening factor until the changes in wage rates, prices, and trade-flow shares are essentially zero, and compute the (final) counterfactual level of welfare, $V_{ij}^c = w_j^c/P_j^c$. From this, we can compute $d\ln V_{ij}$, which equals $(-\gamma)d\ln\lambda_{jj}$. We conduct this process 1358 times for 1358 bilateral EIA removals. Finally, every one of the 1358 simulations yielded unique values for the N national wage rates w_j, supporting section 2's theoretical conjecture of unique wage rates.

We estimate equation (35) using ordinary least squares (OLS). However, as in the gravity equation literature, the relationship between the variables of interest is multiplicative. For OLS, we follow the traditional gravity equation literature – prior to Santos Silva and Tenreyro (2006) – where we assume the error term, ϱ_{ij}, is multiplicative and rewrite equation (35) as:

$$d\ln V_{ij} = -(1/\gamma)\lambda_{ij}\hat{\beta}_{ij}\varrho_{ij} \tag{36}$$

Taking the logarithm of equation (36) yields a log-linear equation suitable for OLS:

$$\ln(d\ln V_{ij}) = \delta_0 + \delta_1 \ln\lambda_{ij} + \delta_2 \ln\hat{\beta}_{ij} + \ln\varrho_{ij}. \tag{37}$$

Our theory suggests the hypothesis that $\delta_1 = \delta_2 = 1$.

Table 10 reports the results of estimating equation (37) under four alternative specifications. Note that, despite having 2460 estimates of $\ln\hat{\beta}_{ij}$, due to the need to construct *intra-national* trade using the World Input-Output Data base for gross output (alongside trade flows), we can only generate 1358 estimates of $\ln\lambda_{ij}$. Specification (1) is equation (37), but constraining the coefficients δ_1 and δ_2 to be equal. Column (3) shows that the coefficient estimate for $\ln(\lambda_{ij}\hat{\beta}_{ij})$ is positive and statistically significant. Moreover, the coefficient estimate of 1.05 is very close to the expected estimate value of 1. Variation in $\ln(\lambda_{ij}\hat{\beta}_{ij})$ explains 98 percent of the variation in $\ln(d\ln V_{ij})$.

Before allowing the coefficients for $\ln\lambda_{ij}$ and $\ln\hat{\beta}_{ij}$ to be unconstrained, specification (2) in column (4) includes just $\ln\lambda_{ij}$; this will allow us to determine shortly the marginal explanatory power of heterogeneous partial EIA effects to explain welfare changes. The R^2 falls substantially from 0.98 to 0.83, and the coefficient estimate for $\ln\lambda_{ij}$ is 1.34.

In specification (3) in column (5), we add $\hat{\beta}_{ij}$ to the regression and allow the coefficient estimates for $\ln\lambda_{ij}$ and $\ln\hat{\beta}_{ij}$ to be unconstrained. Column (5) shows that $\ln\lambda_{ij}$ and $\ln\hat{\beta}_{ij}$ have positive and statistically significant effects on $\ln(d\ln V_{ij})$. Both variables explain 98 percent of the variation in $d\ln V_{ij}$; the addition of $\hat{\beta}_{ij}$ adds 15 percentage points in explanatory power. This suggests that the correlation between the welfare changes and the partial effects is strong. Naturally, we would expect a correlation because the welfare effect is a function of the partial effect; in the next sub-section we will consider for robustness an instrument for the welfare effect used here. Nevertheless, with an explanatory power of 98 percent, our result suggests that general equilibrium factors play a limited role empirically relative to the heterogeneous partial effects in influencing welfare. Note that the coefficient estimate for $\ln\lambda_{ij}$ is 1.02, close to that suggested by theory, and the coefficient estimate for $\ln\hat{\beta}_{ij}$ is 1.09, which is close to unity but is statistically different from unity (at the 1 percent significance level).

Finally, specification (4) in column (6) adds an importer fixed effect to account for general equilibrium effects, which are importer specific, $d\ln w_j - \sum_{k\neq i}^{N}\lambda_{kj}d\ln p_{kj}$. The R^2 value rises from 0.98 to 0.99 with the inclusion of the importer fixed effect. Moreover, while the coefficient estimate for $\ln\lambda_{ij}$ rises slightly to 1.05, the coefficient estimate for $\ln\hat{\beta}_{ij}$ drops from 1.09 to 1.08, with both estimates still close to unity.

On net, the results suggest that welfare changes for importer j from an EIA with exporter i are well-approximated by (partial effect) estimates of $\ln\lambda_{ij}\hat{\beta}_{ij}$. However, since the "data" used for the LHS variable in the regressions just reported ($d\ln V_{ij}$) are generated from a general equilibrium model that incorporates the partial effect estimate, we evaluate next the robustness of these results. We do this by examining the roles of $\ln\hat{\beta}_{ij}$ and $\ln\lambda_{ij}$ for explaining an *empirically generated measure* of the potential welfare gain from an EIA between i and j, suggested by the methodology in Baier and Bergstrand (2004): probit estimates of the likelihood of an EIA.

5.2.2. Robustness analysis

As just noted, one of the constraints of the previous regressions is that the welfare changes are functions of the partial effects by construction. The purpose of the preceding analysis was to show that general equilibrium effects played little role quantitatively. However, there is another way to show that $\ln\hat{\beta}_{ij}$ and $\ln\lambda_{ij}$ are useful and readily available variables for predicting welfare changes from an EIA. Baier and Bergstrand (2004) provided a framework for predicting the probability that a pair of countries would have an EIA. Based upon a general equilibrium model, the authors showed that the welfare of two countries' representative consumers would be enhanced by an EIA the closer they were to each other, the more remote they were from the rest-of-the-world, the larger their economics sizes, and the more similar their economic sizes. Following a qualitative choice model, they showed that these economic factors would also be related to the probability of having an EIA. Their results indicated that the country-pairs that tended to have EIAs tended to have the economic characteristics consistent with such EIAs being welfare improving. Moreover, the econometric model predicted correctly 85 percent of the 286 EIAs in 1996 among the 1431 country-pairs and predicted correctly 97 percent of the remaining 1145 pairs with no EIA.

The econometric framework we employ here is the qualitative choice model, which can be derived from an underlying latent variable model. For instance, let y^* denote an unobserved (or latent) variable, where for simplicity we ignore the observation subscript. As in Wooldridge (2000), let y^* represent the difference in utility levels from an action (the formation of an FTA), where:

Table 10
Determinants of Changes of (Logs of) Welfare.

(1) Variables	(2) Expected Coefficient Value	(3)	(4)	(5)	(6)
$\ln \lambda_{ij}\widehat{\beta}_{ij}$	1	1.046 ***			
		(255.35)			
$\ln \lambda_{ij}$	1		1.337 ***	1.022 ***	1.048 ***
			(80.88)	(158.41)	(150.74)
$\ln \widehat{\beta}_{ij}$	1			1.094 ***	1.081 ***
				(101.27)	(88.72)
Constant	?	3.775 ***	4.561 ***	3.677 ***	5.528 ***
		(132.86)	(46.17)	(105.45)	(23.25)
Fixed Effects:					
Importer		No	No	No	Yes
R^2		0.980	0.828	0.980	0.985
N		1358	1358	1358	1358

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively. t-statistics are in parentheses.

$$y* = \varsigma_0 + x\varsigma + \epsilon \quad (38)$$

where x is a vector of explanatory variables (i.e., common economic characteristics), ς is a vector of parameters, and error term ϵ is assumed to be independent of x and to have a standard normal distribution. In the context of this model, $y^* = \min(\Delta U_i, \Delta U_j)$. Hence, both countries' consumers need to benefit from an EIA for their governments to form one. Since y^* is unobservable, we define an indicator variable, *EIA* which takes the value 1 if two countries have an EIA (indicating $y^* > 0$), and 0 otherwise (indicating $y^* \leq 0$). The response probability, P, for EIA is:

$$P(EIA = 1) = P(y* > 0) = H(\varsigma_0 + x\varsigma) \quad (39)$$

where $H(\cdot)$ denotes the standard normal cumulative distribution function, which ensures that $P(EIA = 1)$ lies between 0 and 1.

In this context, we predicted the probabilities of country-pairs having EIAs for the nine years 1970, 1975, …, 2010 using similar economic characteristics; the probit results are provided in Online Appendix 8. The relationships between the economic characteristics with the probabilities are qualitatively very similar across the nine years and are consistent with findings in Baier and Bergstrand (2004). As expected based upon the theoretical framework in Baier and Bergstrand (2004), the likelihood of an EIA is negatively related to distance, positively related to economic size (joint GDP), and negatively related to GDP dissimilarity. We also find that that the probability of an EIA is positively related to having a common primary language and religious similarity. In the context of the Baier-Bergstrand model, the country-pairs that tend to have EIAs tend to have the economic characteristics consistent with such EIAs being welfare improving.

Our goal in this section is to determine whether $\ln \widehat{\beta}_{ij}$ and $\ln \lambda_{ij}$ can also explain the variation in the probabilities of EIAs, which serve as proxies for the welfare gains of a country-pair from an EIA. One limitation is that, by using a proxy for the welfare change, we *cannot* anticipate coefficient estimates of unity. Table 11 presents the results of five alternative specifications. The number of observations (for year 2005) is limited to 1,358, as these are the number of estimates of $\ln \lambda_{ij}$ from our earlier results. The first specification, shown in column (3), regresses $\ln P(EIA_{ijt})$ on only $\ln \lambda_i$ and a constant. The coefficient estimate for $\ln \lambda_i$ has the expected positive sign. Importantly, note that the pseudo-R^2 value here is only 28 percent.

Specification (2) in column (4) includes both $\ln \lambda_i$ and $\ln \widehat{\beta}_{ij}$. Both coefficient estimates are positive and statistically significant. In this case, the marginal explanatory power from adding $\ln \widehat{\beta}_{ij}$ to the regression is *very large*; the pseudo-R^2 value rises from 28 percent to 75 percent. Recall that, because we are using a proxy for welfare changes, we cannot expect coefficient estimates of unity.

One possible concern, however, is that the $\ln \widehat{\beta}_{ij}$ work well to explain the probability of an EIA because the $\ln \widehat{\beta}_{ij}$ themselves will tend to be higher when variable and fixed non-policy export costs are low, as our theory suggested. Consequently, $\ln \widehat{\beta}_{ij}$ may have an economically and statistically significant effect simply because $\ln \widehat{\beta}_{ij}$ and $\ln P(EIA_{ijt})$ are influenced by common variables, such as bilateral distance, adjacency, etc. To address the robustness of our results, we considered several other specifications. Column (5) adds the log of bilateral distance to the regression. Column (5) shows that – although $DIST_{ij}$ helps to explain $\ln P(EIA_{ijt})$ – the coefficient estimate for $\ln \lambda_{ij}$ becomes negative, but that for $\ln \widehat{\beta}_{ij}$ is still positive and statistically significant. Moreover, adding only distance increases the explanatory power from 75 percent to 91 percent. However, in the next sensitivity analysis, we included bilateral distance as well as all the other variables used earlier to explain variable and fixed export costs (and which are determinants of the predicted probabilities, as shown in Online Appendix 8). Column (6) shows that – although most of these observables are statistically significant in explaining $\ln P(EIA_{ijt})$ – $\ln \widehat{\beta}_{ij}$ *and* $\ln \lambda_{ij}$ now both have positive and statistically significant coefficient estimates again. The specification in column (7) adds an importer fixed effect, as earlier, to account for the importer-specific general equilibrium effects. As shown in column (7), this has a minor effect on the explanatory power of $\ln \lambda_{ij}$ and $\ln \widehat{\beta}_{ij}$. This final specification has a pseudo-R^2 of 95 percent.

6. Relevance of findings to the current trade-policy debate

Starting around 2016, the world has witnessed a rise in protectionism. Some prominent examples in the case of EIAs include "Brexit" (the proposed departure of (technically) the United Kingdom from the European Union) and a renegotiation of the North American Free Trade Agreement (NAFTA). To see the relevance of the findings of this paper for recent efforts to dissolve or dismantle existing EIAs, in this section we conduct three experiments.[44] First, we illustrate in the context of our empirical model the average treatment effect on bilateral trade between the United Kingdom (UK) and other members of the European Union (EU) of the UK withdrawing from the EU (and assuming no offsetting EIA, such as an FTA). Although the referendum vote for Brexit passed in 2016, we consider also two other speculations. Nationalism has been rising in several countries of central and eastern Europe. For purposes of example and in the spirit of this paper's focus on differential effects for developing versus developed economies, we also consider the effects of EU membership for Poland and Croatia – two developing economies in the EU. The results of these three experiments using our methodology

[44] We thank a referee for suggesting this section.

Table 11
Determinants of (Logs of) Probabilities of EIAs.

(1) Variables	(2) Expected Coefficient Sign	(3)	(4)	(5)	(6)	(7)
$\ln \lambda_{ij}$	+	0.765 ***	0.224 ***	−0.090 ***	0.221 ***	0.183 ***
		(23.09)	(9.97)	(-5.67)	(12.11)	(7.89)
$\ln \widehat{\beta}_{ij}$	+		1.877 ***	0.685 ***	0.664 ***	0.743 ***
			(49.95)	(19.14)	(19.19)	(20.24)
$\ln DIST_{ij}$	–			−1.509 ***	−1.471 ***	−1.399 ***
				(-44.71)	(-46.97)	(-38.86)
ADJ_{ij}	?				−1.470 ***	−1.380 ***
					(-22.79)	(-21.54)
$LANG_{ij}$	?				0.218 ***	0.181 ***
					(4.11)	(3.38)
$RELIG_{ij}$	?				−0.254 ***	−0.287 ***
					(-6.80)	(-7.21)
$LEGAL_{ij}$	?				−0.291 ***	−0.202 ***
					(-8.66)	(-6.10)
$COLONY_{ij}$	?				−0.049	−0.030
					(-0.59)	(-0.39)
Constant		2.292 ***	0.774 ***	9.197 ***	10.98 ***	14.18 ***
		(11.58)	(6.38)	(45.18)	(55.13)	(31.59)
Fixed Effects:						
Importer		No	No	No	No	Yes
Pseudo R^2		0.282	0.747	0.898	0.932	0.947
N		1358	1358	1358	1358	1358

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively. t-statistics are in parenthesis.

are provided in Table 12.

First, consistent with the thrust of this paper, there is substantive heterogeneity in the partial ATTs (omitting here analysis of general equilibrium effects). Estimates of the bilateral trade-creation (ATT) impacts for membership of the UK with 25 other (non-Croatia and non-Poland) members of the EU range from a 0.4 percent increase of UK trade with Cyprus (ATT = 0.0037) to a 141 percent increase with the Netherlands (ATT = 0.88). Assuming symmetry, *dis-solution* of the agreement would have symmetric, contractionary effects.

Second, there is also substantive heterogeneity in the ATTs for Croatia and Poland, two countries with much lower per capita GDPs. Moreover, we also chose Croatia and Poland – not just because they are lower per capita GDP countries than the UK but – because their economic centers are approximately the same distance from the economic center of the EU (and we know from our analysis earlier that closer distance increases the partial effect of an EIA). For Croatia, the partial effects (ATTs) from membership of Croatia with 25 other (non-UK and non-Poland) members of the EU range from a 44 per-

Table 12
ATTs for the United Kingdom, Croatia, and Poland.

UNITED KINGDOM			CROATIA			POLAND		
(1) Exporter	(2) Importer	(3) ATT	(4) Exporter	(5) Importer	(6) ATT	(7) Exporter	(8) Importer	(9) ATT
UK	Austria	0.617	Croatia	Austria	0.872	Poland	Austria	0.968
UK	Belgium	0.873	Croatia	Belgium	0.813	Poland	Belgium	0.818
UK	Bulgaria	0.411	Croatia	Bulgaria	0.557	Poland	Bulgaria	0.464
UK	Cyprus	0.004	Croatia	Cyprus	0.467	Poland	Cyprus	0.419
UK	Czech Republic	0.580	Croatia	Czech Republic	0.710	Poland	Czech Republic	0.821
UK	Germany	0.816	Croatia	Germany	0.780	Poland	Germany	0.654
UK	Denmark	0.828	Croatia	Denmark	0.552	Poland	Denmark	0.680
UK	Spain	0.581	Croatia	Spain	0.719	Poland	Spain	0.661
UK	Estonia	0.495	Croatia	Estonia	0.369	Poland	Estonia	0.504
UK	Finland	0.663	Croatia	Finland	0.453	Poland	Finland	0.571
UK	France	0.735	Croatia	France	0.821	Poland	France	0.768
UK	Greece	0.386	Croatia	Greece	0.610	Poland	Greece	0.500
UK	Hungary	0.614	Croatia	Hungary	0.968	Poland	Hungary	0.840
UK	Ireland	0.785	Croatia	Ireland	0.691	Poland	Ireland	0.704
UK	Italy	0.579	Croatia	Italy	0.967	Poland	Italy	0.789
UK	Lithuania	0.549	Croatia	Lithuania	0.675	Poland	Lithuania	0.909
UK	Luxembourg	0.777	Croatia	Luxembourg	0.859	Poland	Luxembourg	0.841
UK	Latvia	0.532	Croatia	Latvia	0.408	Poland	Latvia	0.578
UK	Malta	0.476	Croatia	Malta	0.825	Poland	Malta	0.701
UK	Netherlands	0.880	Croatia	Netherlands	0.697	Poland	Netherlands	0.725
UK	Portugal	0.558	Croatia	Portugal	0.654	Poland	Portugal	0.614
UK	Romania	0.447	Croatia	Romania	0.578	Poland	Romania	0.536
UK	Slovak Republic	0.600	Croatia	Slovak Republic	0.904	Poland	Slovak Republic	0.965
UK	Slovenia	0.581	Croatia	Slovenia	1.115	Poland	Slovenia	0.761
UK	Sweden	0.750	Croatia	Sweden	0.505	Poland	Sweden	0.638

Notes: Bilateral ATT between each country pair is symmetric regardless if country is exporter or importer.

cent increase of Croatian trade with Estonia (ATT = 0.37) to a 203 percent increase with Slovenia (ATT = 1.22). For Poland, the partial effects from membership of Poland with 25 other (non-UK and non-Croatia) members of the EU range from a 52 percent increase of Polish trade with Cyprus (ATT = 0.42) to a 164 percent increase with Austria (ATT = 0.97). Assuming symmetry, *dis-solution* of the agreement would have symmetric, contractionary effects.

Third, closer examination of the results in Table 12 reveals that for 18 of the 25 pairings Croatia and Poland had *larger* partial effects (ATTs) with an EU member than did the UK. Only for pairings of the UK with physically close Ireland, Belgium, the Netherlands, Germany, Denmark, Sweden and Finland were the ATTs larger for the UK than those for either Croatia or Poland. This evidence strongly suggests that developing economies have relatively larger partial effects (and then likely welfare gains) from EIAs than developed economies, and consequently the potential costs of dis-solutions would likely also be larger for developing economies.

7. Conclusions

This paper has offered three contributions. First, extending a standard Melitz model of trade to include additively separable exogenous policy and non-policy fixed export costs along with endogenous fixed export costs (motivated by network effects), we have shown that variable and fixed trade-cost elasticities associated with trade liberalizations are heterogeneous and endogenous to country-pairs' bilateral policy and non-policy, variable and fixed trade-cost *levels* – even allowing for CES preferences and untruncated Pareto productivity distributions. Second, associated comparative statics suggested testable hypotheses for the influence of (observable proxies for) policy and non-policy trade-cost levels on EIA dummy coefficients in a properly specified gravity equation. Panel estimation of the heterogenous partial EIA effects confirmed robustly the expected interactions. Third, we demonstrated the quantitative relevance of these theoretical and empirical results for welfare calculations in the context of the new quantitative trade models.

On March 21, 2018, the heads of 44 of 55 countries in Africa signed the Continental Free Trade Area (CFTA) agreement, in the wake of the effort initiated in 1994 at the Treaty Establishing the African Economic Community (the Abuja Treaty). How is the methodology of this paper potentially relevant? In Economic Commission for Africa (2017), titled *Bringing the Continental Free Trade Area About*, the African Union report's chapter 4 on the economic case for the CFTA notes:

Baier and Bergstrand (2007) *find that,* ***on average****, a free trade area approximately doubles two members' bilateral trade after 10 years …. The ultimate effect of a free trade area depends on the particular characteristics of member countries, including the compatibility of their trade profiles, pre-existing tariff structures, and geographical proximity*. (p. 65; bold added)

While analyzing *ex ante* the trade and welfare effects of this recently signed CFTA is well beyond the scope of this already lengthy paper, even the African Union economists suggest the importance of establishing a methodological framework that extends the homogeneous average partial EIA effects to consider heterogeneous partial EIA effects.

Yet, more work needs to be done. Perhaps the most pressing issue to link the theoretical results to empirical analysis is better data on "trade policies," as emphasized in Goldberg and Pavcnik (2016). Economists agree that *ad valorem* measures of tariff rates are woefully inadequate; measures of policy export fixed costs are virtually non-existent. In this study, the use of dummy variables to capture the treatment effects from EIA liberalizations follows from BB, BBF, and Head and Mayer (2014). However, perhaps the work needed to be done should cast an eye to earlier efforts to measure and analyze the effects of non-tariff measures and trade facilitation for European Union countries addressed in Anderson et al. (2008), which provided underlying methodology for Berden et al. (2010) and Francois et al. (2013). Such improved methodology can likely augment new quantitative trade model estimates of trade liberalizations examined in papers such as Caliendo et al. (2015), which omit policy fixed export cost changes likely associated with EIAs – especially for developing countries.

Appendix A. Supplementary data

Supplementary data related to this article can be found at https://doi.org/10.1016/j.jdeveco.2018.08.014.

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