The University of Nottingham

research paper series

Political Economy of Globalisation

Research Paper 2017/05

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March 17, 2017

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; importing developing countries face higher market-entry costs (partly due to poorer international networks). In this paper, we address how the trade and welfare effects of EIAs are sensitive to the *levels* of country-pairs' variable and fixed trade costs. It is now well established that the (variable-cost) "trade elasticity" – typically estimated using gravity equations of international trade flows – is central to computing general equilibrium impacts on trade and economic welfare of trade-policy liberalizations using the new quantitative trade models. However, this trade elasticity is generally assumed to be an *exogenous parameter*, such as the elasticity of substitution in consumption or an index of heterogeneous productivities; moreover, most studies have ignored the role of the *fixed-export-cost* trade elasticity for trade-policy liberalizations (when not ignored, assumed parametric). This paper offers 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 explicit predictions of the heterogeneous (variable- *and* fixed-cost) bilateral extensive-margin, intensive-margin, and trade elasticities. Second, a modification of the state-of-the-art panel-data methodology to estimate consistent average treatment effects of economic integration agreements (that liberalize variable and fixed trade-policy costs) provides empirical support for the theoretical hypotheses. Consistent with a growing empirical literature, trade elasticities vary across particular settings; such variation is especially pronounced for North-South EIAs. Third, we demonstrate the relevance of these theoretical and empirical results for welfare calculations using the new quantitative trade models. We show empirically that *83-94 percent* of the welfare (or probability) estimates of economic integration agreement liberalizations between 2,266 North-North, North-South, and South-South country-pairs can be explained by our heterogeneous economic integration agreement partial treatment effects.

Key words: International trade, economic integration agreements, gravity equation, welfare
JEL classification: F1, F13, F63, O10, O24

* *Acknowledgements*: To be added.

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"*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)

1 Introduction and Relevant Literature

It is now widely accepted that economic integration agreements (EIAs) and other trade-policy liberalizations contribute to nations' economic 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, Costinot, and Rodriguez-Clare (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 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 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 trade and welfare impacts of trade-policy liberalizations?

To address these questions, this paper offers three potential contributions. First, we extend a

[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, Kaltani, and Loayza (2009). For an overview of how historical factors involving international trade policies matter for economic development, see Nunn (2014).

standard Melitz model of trade with firm heterogeneity and export fixed costs to show theoretically how (variable- *and* fixed-trade-cost) 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 the 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; using simulations they demonstrate that welfare gains can be substantially different depending upon the assumption about the Pareto distribution. Melitz and Redding (2015) also note an emerging empirical literature on heterogeneous trade elasticities. For instance, Helpman, Melitz, and Rubinstein (2008), or HMR, find empirical evidence for endogenous elasticities of trade with respect to distance in the context also 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. 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. However, an endogenous trade elasticity does not rule out CES preferences. 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. Moreover, our model 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. Head and Mayer (2014) used a gravity equation with EIA dummies to determine the welfare gains from EIAs using a new quantitative trade model. Extending here Baier, Bergstrand, and Feng (2014) and Head and Mayer (2014) to estimate consistent treatment effects of EIA dummies, 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."[2] Moreover, due to our accounting properly for endogeneity bias our study provides the clearest evidence to date of the influences of several core bilateral gravity variables in influencing trade elasticities in predictable ways; standard geographic, cultural and institutional variables *all* significantly influence the extensive margin elasticity, whereas only geographic variables (distance and adjacency) influence the intensive margin elasticity.[3]

[2] As noted above, HMR and Novy (2013) only examined heterogeneous trade elasticities to distances.

[3] One of the benefits of using EIA dummies rather than *ad valorem* tariff rates is that the former captures reductions in export fixed costs from EIA formations, whereas tariff rates alone cannot. Baier, Bergstrand, and Feng (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).

Third, 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 2,266 bilateral EIA liberalizations. Consistent with theory, we show that 83-94 percent of the variation in these 2,266 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 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 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.[4] 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 section 3, 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, 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 only distance and adjacency – influencing variable transport costs – explain well the heterogeneity in EIA dummies' partial effects on the *intensive* (product) margin. 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

[4] We will distinguish Krautheim's approach to endogenous fixed export costs from that in Arkolakis (2010) later.

economic variables with EIA dummies, but the study was not guided by theory and so interaction effects lacked economic interpretation. Cheong, Kwak, and Tang (2015) examined empirically interactions of EIA dummies only with measures of GDP size similarity and per capita income similarity and found significant effects. By examining interactions of EIAs dummies with per capita income differences, Cheong, Kwak, and Tang (2015) found preliminary evidence that EIA effects varied by N-N, N-S, and S-S country-pairs. Like Vicard (2011) though, Cheong, Kwak, and Tang (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, Brakman, and Garretsen (2014)).[5] We employ the Hummels and Klenow (2005) product-margin-decomposition methodology, as in Baier, Bergstrand, and Feng (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 nontradable goods' "cutoffs" and interaction effects by type of EIA.[6]

Finally, 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. Studies such as Baier and Bergstrand (2007) and Baier, Bergstrand, and Feng (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 section 4, we show that our method for estimating quantitatively the sensitivity of estimated partial effects to geographic, institutional, and cultural characteristics enables gravity equations to more precisely inform policy makers *ex ante* of *pair-specific* predicted impacts of EIAs – accounting for both heterogeneous general *and partial* equilibrium effects. We 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 83-94 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

[5] 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).

[6] It is important to note that, 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, our theory will be cast with a focus on heterogeneous partial tariff-rate and export-fixed-cost elasticities. We leave for future research applying the methodology in this paper to the case where high quality *ad valorem* measures of bilateral tariffs, non-tariff barriers, and other export fixed costs are available.

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, and we demonstrate this briefly in the context of the proposed Trans-Atlantic Trade and Investment Partnership. Section 5 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) fixed export costs, *and* (additively separable) endogenous fixed export costs (or network effects). In the second part, we solve for a gravity equation analogous to that in Redding (2011). In the third part, we provide comparative statics for *ad valorem* tariff-rate changes that motivate several testable theoretical propositions explored empirically in section 3. In the fourth part, we provide comparative statics for policy *fixed* export-cost changes that motivate several other testable theoretical propositions also explored empirically in section 3.

2.1 The Model

Our theoretical model is an extension of the Redding (2011) version of the Melitz (2003) model. Our model has four distinguishing (and economically plausible) features. 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 why intensive-margin elasticities of bilateral tariff rates are increasing (in absolute value) in bilateral distances between countries. The second is to additively separate exogenous *policy* export fixed costs from exogenous *non-policy* (or natural) export fixed costs; no previous Melitz model has done this. This feature also will be important later for the comparative statics. For instance, we will show theoretically (and later empirically) 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. Our model includes both in an economically plausible way (additively separable), and generates endogenous tariff-rate and policy fixed export cost elasticities. 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 (unlike Krautheim (2012)). This is not a trivial extension; accordingly, Online Appendix 1 develops this extension in a closed-economy Melitz model to prove first the existence and uniqueness 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, i.e., the Redding (2011) version of the Melitz model.[7]

Separating the exogenous component of export fixed costs (A) into two additive parts is straightforward and seemingly trivial. However, as derivations will reveal, this additive separability will make the extensive-margin and trade elasticities of policy fixed export cost changes endogenous to the *levels* of policy and non-policy export fixed costs. Similarly, an additive relationship between gross tariff rates ($t > 1$) and freight rates ($fr > 0$) – the two components of variable trade costs ($\tau > 1$) – can lead to endogenous extensive-margin, intensive-margin, and trade elasticities of tariff-rate changes. Anderson and van Wincoop (2004, p. 715) is the most prominent study to suggest this formulation of the trade-cost factor, $\tau = t + fr$. 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 (1999), 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 $fr_{ij} = freight_{ij}/p_i$.

The notion of endogenous export fixed costs was first introduced in Krautheim (2012), a Chaney (2008) type model with an exogenous number of varieties in each country and no free entry and exit.[8] 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. Yet, in his final section 4, he 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. However, he does not provide a general equilibrium model of these influences in his paper. Moreover, he 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, a notion consistent with the spirit of the quote from Goldberg and Pavcnik (2016).[9]

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

[8] Krautheim (2012) refers to Krautheim (2007) for the microeconomic foundations of his (and hence our) formulation of network influences; we refer the reader to Krautheim (2007) for microeconomic foundations. Krautheim (2012) notes based upon a German chamber of commerce survey that many firms reported obtaining information related to export fixed costs through *indirect* channels involving interactions with business partners and with personal network connections. This is the anecdotal source of network spillovers between exporters. Krautheim (2012) summarizes the microeconomic foundations: "Each firm exporting to country j gains some country-specific knowledge about exporting. The more of this information available to the firm, the cheaper it is to perform export related tasks and therefore the fixed costs of exporting are lower. Firms can then choose to join (costlessly) a network of exporters.... so that in equilibrium all firms join and the fixed costs are lowered for all firms" (p. 33).

[9] Krautheim (2012) notes that the "most related paper" to his is Arkolakis (2010). The latter paper endogenized the fixed export costs of marketing in a foreign market by using marketing technology with decreasing returns that led more productive firms to penetrate more a given market. The Krautheim and Arkolakis approaches are different, but potentially complementary. In the Krautheim model, firms face an industry-wide externality in the origin market;

In this section of the paper, we summarize the more general open-economy model. As in Redding (2011), 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.[10]

Consumers (workers) are identical and have the 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:

$$q_{ij}(\omega) = \left(\frac{p_{ij}(\omega)}{P_j} \right)^{-\sigma} \left(\frac{w_j L_j}{P_j} \right) \tag{2}$$

where $P_j = [\int_{\omega \in \Omega_j} p(\omega)^{1-\sigma} d\omega]^{\frac{1}{1-\sigma}}$, w_j is the wage rate in country j, and $w_j L_j$ is aggregate income in country j which is equal to aggregate expenditure.

Firms are assumed to have heterogeneous productivities. Entry into a market by a firm requires an exogenous cost f_i^e in country i.[11] In order to sell in a market j, a firm has to pay a fixed cost, f_{ij}. We assume furthermore that fixed costs f_{ij} can be decomposed linearly into fixed costs associated with – what we term – "natural" (or non-policy) impediments into markets (such as costs associated with geographic distance or cultural differences) and fixed costs associated with the destination market's trade "policy" impediments (such as costs associated with institutional differences).[12] 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:

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

export fixed costs are endogenous but not a control variable of individual firms. By contrast, Arkolakis (2010) introduces a third ("consumers") margin of adjustment, which is both the focus of the theoretical model as well as the subsequent empirical work focusing on trade growth effects of smaller already exporting firms in a destination market relative to larger already exporting firms. Moreover, market-entry fixed costs in Arkolakis (2010) enter multiplicatively, not additively.

[10]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.

[11]In Redding (2011), f^e is common across countries, but that assumption is unnecessary.

[12]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.

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

Because we are ultimately interested in endogenous trade elasticities from forming EIAs, which lower *ad valorem* tariff rates (and policy-based export fixed costs), it will be useful to separate variable trade costs τ_{ij} into tariff and freight components. As discussed above, assume the *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.[13]

Up to now, our model is standard, except for distinguishing two types of fixed costs (policy and non-policy) and two types of variable costs (t_{ij} and fr_{ij}). We now introduce network effects into the fixed costs f_{ij}. 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) and discussed above, we assume that the fixed costs of selling a product from i to j are inversely related to the 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.[14]

Following Redding (2011), our model now 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) as well as because we have additively separable exogenous and endogenous fixed costs. In this setting, the profits of firm φ in i selling to j (π_{ij}) are:

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

Firms in i will choose to sell to j as long as profits are positive. The marginal exporter from i to j,

[13] 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, Feenstra, Romalis and Taylor (2015, especially Appendix A). Similar to Costinot and Rodriguez-Clare (2014), we introduce tariffs in the variable cost function; however, for simplicity we ignore tariff rebates. As shown in Costinot and Rodriguez-Clare (2014), the model can be easily extended to allow for positive tariff revenue.

[14] 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 and 3, this assumption will ensure in section 4 that welfare can be measured using the standard two sufficient statistics discussed in Arkolakis, Costinot, and Rodriguez-Clare (2012).

where profits approach zero, defines the "cutoff" productivity (φ_{ij}^*):

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

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 and uniqueness of the equilibrium values of the cutoff productivity (φ^*) and average firm profits ($\bar{\pi}$).[15]

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 can 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). Before doing so, 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 $g(\phi) = \gamma\phi^{-(\gamma+1)}$ and the cumulative distribution function is $G(\phi) = 1 - \phi^{-\gamma}$, where we assume $\phi_{min} = 1$ for convenience. Hence, $1 - G(\phi) = \phi^{-\gamma}$.

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

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

where α_i is solved for in Online Appendix 2. We will 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.

We start with zero-profit condition equation (8) and equation (9), defining R_{ij} as variable profits:

$$R_{ij} = \left(\frac{w_i\tau_{ij}}{\rho P_j}\right)^{1-\sigma}\frac{w_jL_j}{\sigma}(\varphi_{ij}^*)^{\sigma-1}$$

and C_{ij} as fixed costs:

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

Since $A_{ij}^N + A_{ij}^P > 0$ (by assumption), there exists a unique and stable cut-off productivity if $\partial C_{ij}/\partial(\varphi_{ij}^*)^{\sigma-1} < \partial R_{ij}/\partial(\varphi_{ij}^*)^{\sigma-1}$ when $C_{ij} = R_{ij}$. This implies:

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

$$\frac{\gamma\eta}{(\sigma-1)} w_j \frac{(\alpha_i L_i)^{-\eta}\left((\varphi_{ij}^*)^{\sigma-1}\right)^{\frac{\gamma\eta}{\sigma-1}}}{(\varphi_{ij}^*)^{\sigma-1}} < \left(\frac{w_i \tau_{ij}}{\rho P_j}\right)^{1-\sigma} \frac{w_j L_j}{\sigma}(1+\theta_{ij})$$
$$= (\varphi_{ij}^*)^{1-\sigma} w_j \left[A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}\left((\varphi_{ij}^*)^{\sigma-1}\right)^{\frac{\eta\gamma}{\sigma-1}}\right](1+\theta_{ij}) \tag{10}$$

where:

$$\theta_{ij} = \frac{(\varphi_{ij}^*)^{-\gamma}\alpha_i L_i \left(\frac{w_i \tau_{ij}}{\varphi_{ij}^*}\right)^{1-\sigma}}{\sum_i (\varphi_{ij}^*)^{-\gamma}\alpha_i L_i \left(\frac{w_i \tau_{ij}}{\varphi_{ij}^*}\right)^{1-\sigma}}. \tag{11}$$

Hence, θ_{ij} reflects the relative importance of j's purchases from i in j's total expenditures.

Equation (10) simplifies to:

$$\frac{\gamma\eta}{\sigma-1} \frac{(\alpha_i L_i)_i^{-\eta}\left((\varphi_{ij}^*)^{\sigma-1}\right)^{\frac{\gamma\eta}{\sigma-1}}}{\left[A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}((\varphi_{ij}^*)^{\sigma-1})^{\frac{\eta\gamma}{\sigma-1}}\right]} \frac{1}{(1+\theta_{ij})} < 1.$$

If we define s_{ij} as the relative importance of bilateral endogenous fixed costs in bilateral total fixed costs:

$$s_{ij} = \frac{(\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\gamma\eta}}{A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\gamma\eta}} = \frac{M_{ij}^{-\eta}}{A_{ij}^N + A_{ij}^P + M_{ij}^{-\eta}} = \frac{1}{1 + \frac{A_{ij}^N + A_{ij}^P}{M_{ij}^{-\eta}}} \tag{12}$$

the stability condition reduces to:

$$\frac{\gamma}{\sigma-1}\eta s_{ij}\frac{1}{(1+\theta_{ij})} < 1. \tag{13}$$

The complete set of derivations is in Online Appendix 2. In the limit, as φ^* approaches 0, θ_{ij} approaches 1, but s_{ij} goes to 0; hence, the condition is satisfied. In the limit, as φ^* approaches ∞, θ_{ij} approaches 0, but s_{ij} goes to 1. Hence, a sufficient condition for stability is $\frac{\gamma\eta}{\sigma-1} < 1$. Krautheim (2012) assumed the same condition for an interior solution; in his model, s_{ij} is assumed to be 1 and $(1+\theta_{ij})^{-1}$ is assumed 1. Common to such Melitz models, $\gamma/(\sigma-1)$ is assumed to exceed unity; hence, $\eta < 1$. However, in our model, the stability condition is even more likely to hold than that in Krautheim (2012) since $s_{ij} < 1$ and $(1+\theta_{ij})^{-1} < 1$. Finally, note that the sufficient condition here is analogous to that in Online Appendix 1 for the closed-economy case, $[\gamma/(\sigma-1)]\eta < 1$.

In Online Appendix 2, we show that average firm profits $\bar{\pi}$ are the standard condition:

$$\bar{\pi}_i = w_i f_i^e (\varphi_{ii}^*)^\gamma \ . \tag{14}$$

We now have N free-entry conditions and N^2 ZPC conditions. In Online Appendix 2, we solve for N endogenous $P_i^{1-\sigma}$, N^2 endogenous M_{ij}, N^2 endogenous trade flows X_{ij}, and N (implicit) endogenous wage rates w_i.[16]

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:

$$X_{ij} = \underbrace{\left[\frac{1-G(\varphi_{ij}^*)}{1-G(\varphi_{ii}^*)}\right] M_i}_{\text{Extensive}} \int_{\varphi_{ij}^*}^{\infty} \left(\frac{w_i \tau_{ij}}{\rho \varphi P_j}\right)^{1-\sigma} w_j L_j \frac{g(\varphi)}{1-G(\varphi_{ij}^*)} d\varphi \tag{15}$$

Using the Pareto distribution $g(\varphi) = \gamma \varphi^{-(\gamma+1)}$, $1-G(\varphi_{ij}^*) = (\varphi_{ij}^*)^{-\gamma}$, and $1-G(\varphi_{ii}^*) = (\varphi_{ii}^*)^{-\gamma}$, and that $M_{ij} = \left[[1-G(\varphi_{ij}^*)]/[1-G(\varphi_{ii}^*)]\right] M_i$, then[17]

$$\begin{aligned} X_{ij} =& M_{ij} \int_{\varphi_{ij}^*}^{\infty} \left(\frac{w_i \tau_{ij}}{\rho \varphi P_j}\right)^{1-\sigma} \frac{w_j L_j}{\sigma} \sigma \gamma \varphi^{-(\gamma+1)} (\varphi_{ij}^*)^{\gamma} \\ =& M_{ij} \left(\frac{w_i \tau_{ij}}{\rho \varphi P_j}\right)^{1-\sigma} \frac{w_j L_j}{\sigma} \sigma \gamma (\varphi_{ij}^*)^{\gamma} \int_{\varphi_{ij}^*}^{\infty} \varphi^{-\gamma+\sigma-2} d\varphi \end{aligned}$$

Using equation (9), and solving the integral yields:

$$X_{ij} = (\alpha_i L_i)(\varphi_{ij}^*)^{-\gamma} \left(\frac{\sigma\gamma}{\gamma-(\sigma-1)}\right) (\varphi_{ij}^*)^{\sigma-1} \left(\frac{w_i \tau_{ij}}{\rho P_j}\right)^{1-\sigma} \left(\frac{w_j L_j}{\sigma}\right)$$

where $\alpha_i = (\sigma-1)/(\gamma \sigma f_i^e)$.

Using equation (8):

$$X_{ij} = (\alpha_i L_i)(\varphi_{ij}^*)^{-\gamma} \left(\frac{\sigma\gamma}{\gamma-(\sigma-1)}\right) w_j A_{ij} \left[1 + \frac{(\alpha_i L_i)^{-\eta} (\varphi_{ij}^*)^{\eta\gamma}}{A_{ij}}\right]. \tag{16}$$

Equation (16) 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). 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*

[16] Note that if we assume no network externalities, i.e., $\eta = 0$, then our model simplifies to the same model as in Redding (2011), where wage rates are determined explicitly.

[17] Recall, $\varphi_{min} = 1$ by assumption, for simplicity of notation.

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). Moreover, 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 our empirical specifications in section 3.

Finally, it will be useful to write equation (16) above also as:

$$X_{ij} = \left[\left(\frac{\sigma\gamma}{\gamma-(\sigma-1)}\right)(\alpha_i L_i)\right][w_j]\left[(\varphi_{ij}^*)^{-\gamma}A_{ij}\left(1+\frac{(\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\eta\gamma}}{A_{ij}}\right)\right]. \tag{17}$$

For the remainder of the paper, the first RHS term in brackets is exogenous and unchanged. The third RHS term in brackets represents *partial* effects on extensive margins and trade flows of trade-policy changes from changes in τ_{ij} or A_{ij} *holding constant* income effects via changes in w_j, as discussed below in sections 2.3 and 2.4. Finally, changes in $[w_j]\left[(\varphi_{ij}^*)^{-\gamma}A_{ij}\left(1+\frac{(\alpha_i L_i)^{-\eta}(\varphi_{ij}^*)^{\eta\gamma}}{A_{ij}}\right)\right]$ will be referred to as *general equilibrium* effects of changes in τ_{ij} or A_{ij} (allowing wage rates w_j to change also), which we will quantify later in section 4 when we discuss welfare effects.

2.3 Comparative Statics for *Ad Valorem* Tariff Rates

The introduction of additively separable exogenous and endogenous (export) fixed costs into a standard Melitz model with N countries yields several new insights, which could not be generated in Redding (2011) (which assumed only exogenous fixed costs) nor in Krautheim (2012) (which assumed only endogenous fixed costs). In Krautheim (2012), the introduction of endogenous fixed costs without (independent) exogenous ones generated a "magnifying" of the trade-cost elasticity, but not an "endogenizing" of it. In this section and the next, we use partial equilibrium comparative statics to illustrate the 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 4.[18]

For tractability, in this section we examine three comparative statics from the model; comparative statics 1-3 are related to an exogenous 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 ($d\ln A_{ij}^P$). All comparative statics derivations are in Online Appendix 3 (and allow the multilateral price term, P_j, to change).

[18]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.

2.3.1 Comparative Static 1: Extensive Margin

We can decompose the change in the aggregate trade flow into changes in the intensive and extensive margins. Aggregate trade can be written as:

$$X_{ij} = w_i L_i \int_{\varphi_{ij}^*}^{\infty} x_{ij}(\varphi) dG(\varphi)$$

Using Leibniz rule to separate the intensive and extensive margins, differentiation with respect to τ_{ij} yields:

$$dX_{ij} = \left[w_i L_i \int_{\varphi_{ij}^*}^{\infty} \frac{\partial x_{ij}(\varphi)}{\partial \tau_{ij}} dG(\varphi) \right] d\tau_{ij} - \left[w_i L_i x(\varphi_{ij}^*) G'(\varphi_{ij}^*) \frac{\partial \varphi_{ij}^*}{\partial \tau_{ij}} \right] d\tau_{ij} \tag{18}$$

The first RHS term is the intensive margin change and the second RHS term is the "extensive" margin change, for which the latter is now defined to include the "composition" change, cf., Head and Mayer (2014).

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:

$$\frac{d \ln EM_{ij}}{d \ln t_{ij}} = -\left(\frac{1}{1 + \frac{fr_{ij}}{t_{ij}}} \right) \left(\frac{\gamma - (\sigma - 1)}{1 - \frac{\gamma}{\sigma - 1} \eta s_{ij}} \right) < 0 \tag{19}$$

where s_{ij} and exogeneous export fixed costs ($A_{ij}^N + A_{ij}^P$) are inversely related. Several key insights are revealed by equation (19). 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_{ij}^N) or policy fixed export costs (A_{ij}^P), 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:

$$\frac{d\ln IM_{ij}}{d\ln t_{ij}} = -\left(\frac{1}{1+\frac{fr_{ij}}{t_{ij}}}\right)(\sigma-1) < 0 \tag{20}$$

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} implied by the more economically plausible trade-cost function recommended in Anderson and van Wincoop (2004) 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).

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:

$$\frac{d\ln X_{ij}}{d\ln t_{ij}} = -\left(\frac{1}{1+\frac{fr_{ij}}{t_{ij}}}\right)\left[(\sigma-1)+\left(\frac{\gamma-(\sigma-1)}{1-\frac{\gamma}{\sigma-1}\eta s_{ij}}\right)\right] < 0 \tag{21}$$

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. The second and third lines of the top panel in Table 1 summarize the qualitative effects of lower natural and policy export fixed costs, 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.[19]

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}.[20]

[19] 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).

[20] We do not address in this paper relationships among tariff-rate, transport-cost, and export fixed cost elasticities recently explored in Besedes and Cole (2017). For now, we leave these issues (including the capture of tariff revenue) for subsequent research.

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 (A_{ij}^P). It is important to note that this is the first study to determine potential endogeneity and heterogeneity of *bilateral export fixed cost elasticities* of extensive margins. Previous studies such as Novy (2013) and Melitz and Redding (2015) have only looked at endogenous *ad valorem* variable-trade-cost trade elasticities and Krautheim (2012) only examined *ad valorem* variable-trade-cost trade elasticities.

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:

$$\frac{d\ln EM_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\frac{\gamma}{\sigma-1}-1}{1-\frac{\gamma}{\sigma-1}\eta s_{ij}}\right)\left(\frac{A_{ij}^P}{A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}(\phi_{ij}^*)^{\gamma\eta}}\right) < 0. \tag{22}$$

There are two important insights to glean from equation (22). First, equation (22) 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 (as well as τ_{ij}) 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 equation (22). 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 (22) 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 (22), 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).

As for variable-trade-cost extensive margin elasticities, *even if network externalities did not exist*, lower initial exogenous policy fixed export costs A_{ij}^P could lead to endogenous extensive margin elasticities. When $\eta = 0$:

$$\frac{d\ln EM_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\gamma}{\sigma-1} - 1\right)\left(\frac{A_{ij}^P}{A_{ij}^P + A_{ij}^N}\right) < 0 \tag{23}$$

implying the elasticity of EM_{ij} with respect to A_{ij}^P is sensitive to the relative levels of A_{ij}^N and A_{ij}^P. Moreover, with $\eta = 0$, a lower level of A_{ij}^P will unambiguously cause the extensive margin elasticity to a one percent change in A_{ij}^P to decline. Thus, endogenous export fixed costs elasticities in this paper surface with CES preferences, with an untruncated Pareto productivity distribution, and with *or without* network externalities.

Finally, note in equation (23) that if we assume no natural export fixed costs ($A_{ij}^N = 0$) and no network externality ($\eta = 0$) as in Chaney (2008), equation (23) simplifies to:

$$\frac{d\ln EM_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\gamma}{\sigma-1} - 1\right) < 0 \tag{24}$$

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:

$$\frac{d\ln IM_{ij}}{d\ln A_{ij}^P} = 0 \tag{25}$$

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:

$$\frac{d\ln X_{ij}}{d\ln A_{ij}^P} = -\left(\frac{\frac{\gamma}{\sigma-1} - 1}{1 - \frac{\gamma}{\sigma-1}\eta s_{ij}}\right)\left(\frac{A_{ij}^P}{A_{ij}^N + A_{ij}^P + (\alpha_i L_i)^{-\eta}(\phi_{ij}^*)^{\gamma\eta}}\right) < 0 \tag{26}$$

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, Data, and Empirical Results

In the first section, we discuss the econometric model. 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. 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 regression specifications. In the fifth section, we present the empirical results, which include several robustness analyses.

3.1 Econometric Approach

Many of the trade-policy liberalizations in the past 25 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 fixed export costs.[21] For instance, see Horn, Mavroidis, and Sapir (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.

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 the "average treatment effects" of EIAs on trade flows, cf., Baier and Bergstrand (2007), or BB, Anderson and Yotov (2011), Eicher, Henn, and Papageorgiou (2012), and Head and Mayer (2014). For instance, BB showed that consistent and precise estimates of EIAs on bilateral trade flows could be captured using the gravity-equation specification below using ordinary least squares (OLS):[22]

$$\ln X_{ijt} = \alpha + \eta_{it} + \theta_{jt} + \psi_{ij} + \beta EIA_{ijt} + \upsilon_{ijt} \tag{27}$$

where η_{it} is an exporter-year fixed effect, θ_{jt} is an importer-year fixed effect, ψ_{ij} is a pair fixed effect, and υ_{ijt} is an error term. Equation (27) 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

[21] 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.

[22] For now, we ignore zero trade flows, allowing a log-linear gravity equation; we address this below. See BB and Baier, Bergstrand, and Feng (2014) for theoretical gravity-equation motivation for equation (27).

fixed effect for controlling for the endogeneity of the EIA variable, alongside fixed effects η_{it} and θ_{jt} to account for exporters' and importers' time-varying GDPs and multilateral price terms.

There are several limitations to specification (27), many of which were addressed subsequently in Baier, Bergstrand, and Feng (2014) and Bergstrand, Larch, and Yotov (2015). One limitation of equation (27) 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, Bergstrand, and Feng (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:

$$\begin{aligned}\ln X_{ijt} = \alpha + \eta_{it} + \theta_{jt} + \psi_{ij} + \beta_1 OWPTA_{ijt} + \beta_2 TWPTA_{ijt} + \beta_3 FTA_{ijt} \\ + \beta_4 CUCMECU_{ijt} + \upsilon_{ijt}\end{aligned} \tag{28}$$

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 (27) (or (28)) 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 (27) or (28), this heterogeneity in EIAs' partial effects is captured in the error term, υ_{ijt}, which is *assumed* to be uncorrelated with the other right-hand-side (RHS) variables. 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:

$$\ln X_{ijt} = \alpha + \eta_{it} + \theta_{jt} + \psi_{ij} + \beta_{ij} EIA_{ijt} + \upsilon_{ijt} \tag{29}$$

where the partial effect of an EIA on $\ln X_{ijt}$ is allowed to be *pair-specific*. Theoretical section 2 suggests there exists a set of variables Z_{ij} such that:

$$E(\ln X_{ijt} \mid \alpha, \eta_{it}, \theta_{jt}, \psi_{ij}, \beta_{ij}, EIA_{ijt}, Z_{ij}) = \alpha + \eta_{it} + \theta_{jt} + \psi_{ij} + \beta_{ij} EIA_{ijt} \tag{30}$$

[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.

Without knowing the true values of the β_{ij}, we take expectations over all variables to obtain:

$$E(\ln X_{ijt} \mid \alpha, \eta_{it}, \theta_{jt}, \psi_{ij}, EIA_{ijt}, Z_{ij}) = \alpha + \eta_{it} + \theta_{jt} + \psi_{ij}$$

$$+E(\beta_{ij} \mid \alpha, \eta_{it}, \theta_{jt}, \psi_{ij}, EIA_{ijt}, Z_{ij})EIA_{ijt}$$

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

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

where $Z_{ij} - \mu_Z$ denotes the de-meaned values of Z_{ij}. Absent knowledge of β_{ij}, following Cameron and Trivedi (2005) we should estimate instead:

$$E(\ln X_{ijt} \mid \alpha, \eta_{it}, \theta_{jt}, \psi_{ij}, EIA_{ijt}, Z_{ij}) = \alpha + \eta_{it} + \theta_{jt} + \psi_{ij} + \beta EIA_{ijt} + b_Z(Z_{ij} - \mu_Z)EIA_{ijt} \quad (31)$$

The main goal of this empirical section of the paper is to identify the variables in Z_{ij} to determine the best linear unbiased predictors.

While incorporating theory-motivated interaction terms is a distinguishing feature of the empirical work in this paper, we will also acknowledge (in several sensitivity analyses and caveats) several remaining shortcomings in specifications (27), (28), and (31) that have surfaced in the literature. A third limitation of specification (27), (28), and (31) is the absence of lags. However, BB, BBF, Bergstrand, Larch, and Yotov (2015) (or BLY), and this paper all have sensitivity analyses for lagged EIA influences.

A fourth potential limitation of specifications (27), (28), and (31) is that they do not allow for unobserved *time-varying* effects of falling transport costs and other fixed export costs that have increased international trade over time. To address this issue, BBF also provided estimates using a first-difference specification – that then would typically lead researchers to omit pair fixed effects – but included pair fixed effects. This "random growth first difference" (RGFD) specification (suggested in Wooldridge (2000)) allowed for time-varying pair-specific unobservables.[25] Yet, the different specifications yielded similar findings. Similarly, BLY allowed for unobservable time-varying pair-specific sources of changes in international relative to intranational trade (such as unobservable falling bilateral export fixed costs) by including a dummy variable for international relative to intranational trade interacted with year dummies. Using their OLS specification for a twenty year period, pairings of 89 countries, and aggregate trade flows, BLY found no material difference between the total estimated partial EIA effect with or without the extra terms (see BLY's Table 5). Consequently, this potential limitation is not empirically material.

A fifth potential limitation of specifications (27), (28), and (31) is accounting for zeros in trade. Ignoring zeros could potentially bias results, due to country selection; moreover, one must account for potential bias due to firm heterogeneity in aggregate data, cf., HMR. However, BBF showed

[25]The Online Appendix to BBF provides the first-difference as well as fixed effects supplementary results; see http://www.sciencedirect.com/science/article/pii/S0022199614000506 .

that potential bias due to country selection and firm heterogeneity was largely cross sectional in nature and could be accounted for in panel data by the pair fixed effects; see BBF and its Online Theoretical Supplement for a comprehensive discussion.[26]

Santos Silva and Tenreyro (2006) first discussed zeros in trade and heteroskedasticity (due to Jensen's inequality), and proposed the Poisson Quasi-Maximum Likelihood (PQML) estimator as an alternative. A large literature on PQML and alternative estimators has surfaced to address these issues, which will not be resolved here; see Head and Mayer (2014) for an excellent recent discussion. One other limitation of using the various nonlinear estimators Head and Mayer (2014) discuss is that the relationship between coefficients for trade, extensive margin, and intensive margin effects will not satisfy the Hummels-Klenow adding-up relationships, which will be addressed below.[27]

A final issue is that previous estimates of EIA partial effects have been *ex post*. Historically, the link between estimated *ex post* gravity EIA partial effects and the *ex ante* welfare effects of a trade liberalization has been – at best – tenuous. In this paper, our identification of the geographic, institutional and cultural factors that explain heterogeneous EIA partial effects β_{ij} can be used potentially for predicting *ex ante* the partial trade effect of a specific country-pair's EIA. We show later in section 4 that the change in welfare in importing country j from an EIA with exporting country i can be approximated by the product of β_{ij} (using estimates from b_Z), the share of i's exports in j's aggregate expenditures (λ_{ij}), an estimate of the Pareto shape parameter γ, and an error term capturing remaining general equilibrium influences.

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_{ij}^N), and exogenous policy fixed export costs (A_{ij}^P) discussed in section 2? Beginning with Tinbergen (1962), the empirical gravity equation literature provides 50 years of econometric examination of observable bilateral variables that likely

[26]Briefly, in their online supplementary robustness analysis cited in the previous footnote, BBF estimated probit regressions of the probability of positive trade similar to HMR, but separately for all eight years of their sample (1965, 1970, ..., 2000). Following HMR, BBF used the predicted probit probabilities to construct the $\hat{\bar{\eta}}*_{ijt}$, $\hat{\bar{z}}*_{ijt}$, $\hat{\bar{z}}*_{ijt}^2$, $\hat{\bar{z}}*_{ijt}^3$, and W_{ijt} (which is the second-stage control variable for the fraction of firms (possibly zero)). For the RGFD specifications, BBF used the first-differenced values as controls for $\Delta \ln W_{ijt}$. The results for the RGFD second-stage regressions were reported in Appendix Table A4 in their Online Supplement. We draw attention to two notable results. First, we note that the coefficient estimates for $\hat{\bar{\eta}}*_{ijt}$, $\hat{\bar{z}}*_{ijt}$, $\hat{\bar{z}}*_{ijt}^2$, and $\hat{\bar{z}}*_{ijt}^3$ were qualitatively identical to those in HMR but were *only* statistically significant for aggregate trade flows and for the Hummels-Klenow extensive margin. This accords with the theoretical conjecture based upon the theoretical HMR framework in terms of V_{ijt}; the influence of V_{ijt} (and hence W_{ijt}) works primarily on aggregate trade via the extensive margin. Second, a comparison of Set 1 (Set 2) in BBF Online Appendix Table A4 with the corresponding results in Set 1 (Set 2) in Table 1 in the BBF reveals that the results for the four EIA variables were identical qualitatively and quite similar *quantitatively*. This is in contrast to the findings in HMR for a single cross-section. The reason is that – in the panel specifications – the first-differencing of the data has controlled for the cross-sectional variation in $\hat{\bar{\eta}}*_{ijt}$ and the factors influencing W_{ijt} and the inclusion of ij fixed effects has removed any slow-moving (trend) variation in $\hat{\bar{\eta}}*_{ijt}$ and the factors influencing W_{ijt}.

[27]Also, due to our specification using a very large number of fixed effects, researchers have only been able to obtain convergence under PQML for a limited time series in the panel (i.e., a short panel), cf., Bergstrand, Larch, and Yotov (2015). Consequently, due to our long panel, this further limits us to only use OLS. We also note that Bergstrand, Larch, and Yotov (2015) found, if anything, that OLS biased downward the EIA partial effect estimates relative to PQML estimates.

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, Larch, Staub, and Winkelmann (2011) (or ELSW) have shown empirically that 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, Sapienza, and Zingales (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 (absolute) extensive margin EIA elasticity and trade-flow EIA elasticity, but have no impact on the intensive margin EIA elasticity.

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 the 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, all bilateral variables are from CEPII.[28]

3.3 Other Data

The only 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/.[29] 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. One of the strengths of the Baier-Bergstrand EIA panel is, for 98.6 percent of the cells where the EIA status of the country pair changes (from 0 to 1, 0 to 3, 2 to 3, etc.), there exists a hyperlink to a copy (PDF format) of the original treaty. There are several versions of the data set; the one used for this paper is a 2014 cleaned, extended-to-2011 version of the May 2013 data set. 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

[28] See $http://www.cepii.fr/cepii/en/bdd_modele/presentation.asp?id = 6$ and $http://www.cepii.fr/cepii/en/bdd_modele/presentation.asp?id = 8$.

[29] This data set was constructed under National Science Foundation grants SES-0351018 and SES-0351154 and includes annually from 1960-2011 for the pairings of 195 countries an index ranging from 0-6 of the level of any EIA between the pair.

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.[30] 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 available disaggregate trade data.[31] 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:

$$EM_{ijt} = \frac{\sum_{m \in M_{ijt}} X_{Wjt}^m}{\sum_{m \in M_{Wjt}} X_{Wjt}^m} \tag{32}$$

where X_{Wjt}^m 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:

$$IM_{ijt} = \frac{\sum_{m \in M_{ijt}} X_{ijt}^m}{\sum_{m \in M_{ijt}} X_{Wjt}^m} \tag{33}$$

where X_{ijt}^m 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

[30] Due to space constraints here, see BBF, p. 342 and the BBF Online Appendix.

[31] 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, Kortum, and Kramarz (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, Demidova, Klenow, and Rodriguez-Clare (2008).

two margins equals the ratio of exports from i to j relative to country j total imports:

$$EM_{ijt}IM_{ijt} = \frac{\sum_{m \in M_{ijt}} X_{ijmt}}{\sum_{m \in M_{Wjt}} X_{Wjmt}} = X_{ijt}/X_{jt} \tag{34}$$

where X_{jt} denotes j's imports from the world. Taking the natural logs of equation (34) and some algebra yields:

$$\ln X_{ijt} = \ln EM_{ijt} + \ln IM_{ijt} + \ln X_{jt}. \tag{35}$$

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. We note four issues regarding the HK methodology. First, the term $\ln X_{jt}$ will be subsumed in an importer-time fixed effect. Second, HK applied their methodology to only a cross section. By contrast, we are applying it to a time series of cross sections. Consequently, the trade weights used in constructing EM_{ijt} and IM_{ijt} will likely vary from year to year. To address this, BBF also considered in a sensitivity analysis a chain-weighting technique. However, their results were robust to this alternative weighting procedure. Third, there are numerous zeros in the variables and the results may be biased by ignoring the existence of firm heterogeneity. However, as discussed earlier (and in BBF and the BBF Online Appendix), our panel approach largely alleviates sample-selection bias and firm-heterogeneity bias as raised in HMR. Fourth, the trade data are 5-digit SITC (Revision 1) data from COMTRADE for the period 1962-2010. This is the most disaggregated publicly available data set for bilateral trade flows for a large number of years and a large number of country pairs, constructed on a consistent basis, necessary for the analysis at hand. The 5-digit SITC level data is a higher level of disaggregation than the 4-digit SITC data used in Hillberry and McDaniel (2002), Kehoe and Ruhl (2009), and Foster, Poeschl, and Stehrer (2011).[32]

3.4 Specifications

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

$$\begin{aligned} \ln EM_{ijt} &= \alpha_0 + \eta_{it} + \theta_{jt} + \psi_{ij} + \alpha_1 EIA_{ijt} + \alpha_2(EIA_{ijt} * \ln DIST_{ij}) \\ &+ \alpha_3(EIA_{ijt} * ADJ_{ij}) + \alpha_4(EIA_{ijt} * LANG_{ij}) + \alpha_5(EIA_{ijt} * RELIG_{ij}) \\ &+ \alpha_6(EIA_{ijt} * LEGAL_{ij}) + \alpha_7(EIA_{ijt} * COLONY_{ij}) + \zeta_{ijt} \end{aligned} \tag{36}$$

$$\begin{aligned} \ln IM_{ijt} &= \phi_0 + \eta_{it} + \theta_{jt} + \psi_{ij} + \phi_1 EIA_{ijt} + \phi_2(EIA_{ijt} * \ln DIST_{ij}) \\ &+ \phi_3(EIA_{ijt} * ADJ_{ij}) + \phi_4(EIA_{ijt} * LANG_{ij}) + \phi_5(EIA_{ijt} * RELIG_{ij}) \\ &+ \phi_6(EIA_{ijt} * LEGAL_{ij}) + \phi_7(EIA_{ijt} * COLONY_{ij}) + \nu_{ijt} \end{aligned} \tag{37}$$

[32]See BBF, p. 344 for a comprehensive discussion of the need for and adequacy of 5-digit level disaggregation for studies of this type using a very large number of countries and very long time series in the panel.

$$\begin{aligned}\ln X_{ijt} &= \beta_0 + \eta_{it} + \theta_{jt} + \psi_{ij} + \beta_1 EIA_{ijt} + \beta_2(EIA_{ijt} * \ln DIST_{ij}) \\ &+ \beta_3(EIA_{ijt} * ADJ_{ij}) + \beta_4(EIA_{ijt} * LANG_{ij}) + \beta_5(EIA_{ijt} * RELIG_{ij}) \\ &+ \beta_6(EIA_{ijt} * LEGAL_{ij}) + \beta_7(EIA_{ijt} * COLONY_{ij}) + \upsilon_{ijt} \qquad (38)\end{aligned}$$

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, and $COLONY_{ij}$ is a dummy assuming the value 1 if i and j share a common colonial history and 0 otherwise. 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 PQML. Table 2 summarizes the expected coefficient signs for each of these specifications.

3.5 Empirical Results

3.5.1 Main Empirical Results with EIA Interactions

Table 4 provides the results of estimating equations (36), (37), and (38) using OLS and panel data for every five years from 1965-2010. Columns (2), (4) and (6) provide the expected coefficient signs for the variables' coefficients for the extensive margin, intensive margin, and trade flows 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.

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 *all 16 coefficient estimates* with definitive predictions have the expected coefficient sign.

Second, note that the coefficient estimates for EIA are positive in all three columns. Hence, at the mean 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 extensive margin. Distance and a common land border have negative effects on the (absolute value of the) extensive margin EIA elasticity. Cultural variables common language and religious similarity have positive effects on the (absolute) extensive margin elasticity. 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^N_{ij}), 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^P_{ij}), lowering the extensive margin EIA elasticity. The latter results suggest that the reduction in export fixed costs from an EIA may dominate 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.

However, *as predicted* 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.

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

Overall, the results strongly support the model's predictions and comparative statics. Moreover, Figure 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 main Melitz model.

In the next three sections, we pursue sensitivity analyses. Our robustness analyses address three issues for which the results may be 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.[33]

3.5.2 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, 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 5, for instance, provides the results using the 100,000 US dollar 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 5 religion has instead a significant impact on EIA's intensive margin effect.

[33] As discussed earlier in section 3.1, based upon sensitivity analyses in BBF (and reported in that paper's Online Appendix), we do not provide here any robustness analyses using the BBF random growth first difference (RGFD) specifications or the HMR technique for firm heterogeneity and zeros in trade flows. As discussed earlier, a RGFD specification including pair fixed effects can account for unobserved *time-varying* trade-cost changes; however, the RGFD specifications in BBF yielded EIA effects similar to those using the fixed effects (FE) specifications. Also as shown in BBF's Online Appendix, the second-stage panel regressions from using the HMR approach yielded similar results to the FE specifications. Nevertheless, such robustness analyses could be performed upon request.

3.5.3 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 6 and 7. Table 6 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. Consistent with Table 4, 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. 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 measurable effects on the intensive margin elasticities and have significant negative effects on the extensive margin elasticities, as before.

Table 7 provides the results for customs unions extracted from Online Appendix 5, Table 1; again, there are 21 coefficient estimates (and t-statistics). Consistent with Tables 4 and 6, 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.

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 Figures 2-5 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.

3.5.4 Lagged Effects

BB and BBF estimated treatment effects also allowing for lags. We augmented the model using equation (38) for aggregate trade flows to include five-year lags of the RHS variables, which reduced the sample size. Basically, the results are largely insensitive to including lags. For brevity of space, the results for aggregate trade flows are presented in Table 3 of Online Appendix 5.

4 Welfare

"*The welfare effects in this class of [quantitative trade] models are linked to the change in the share of trade that takes place inside a country.... Intuitively, because the initial flows are so small, even doubling trade with ex-colonies will result in very tiny changes in the share of expenditure that is spent locally. In contrast, adding even a few percentage points of trade with a major partner will be much more important for welfare.*" (Head and Mayer (2014), p. 170)

Up to now, we have focused on the notion that partial trade-cost elasticities may be endogenous and we have provided empirical support using EIA dummies that they are. As an important consideration is the general equilibrium welfare gains from a trade-policy liberalization, it is reasonable to consider – as suggested by the quote from Head and Mayer (2014) above – how important *quantitatively* such endogenous partial trade impacts are for overall welfare gains. This is the issue examined in this section.

This section has four parts. First, following Redding (2011), we show theoretically that welfare in our model can be measured using the same two sufficient statistics as in Costinot and Rodriguez-Clare (2014), after adding one simplifying assumption. Second, 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 overwhelming explained by these two terms, which may be much easier to measure for a large number of countries and large number of EIAs. Third, 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. Fourth, we discuss briefly the implications for an *ex ante* analysis of the proposed Trans-Atlantic Trade and Investment Partnership.

4.1 Measuring Welfare

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, Costinot, and Rodriguez-Clare (2012).

As shown in Online Appendix 6, using the ZPC condition and the trade-share equations, we can show that indirect welfare in our model (V_j) can be expressed as:

$$V_j = w_j/P_j = \lambda_{jj}^{-\frac{1}{\gamma}} L_j^{\frac{1}{\sigma-1}} \left[\frac{A_{jj}^{1-\frac{\gamma}{\sigma-1}}}{f_j^e (\frac{\sigma}{\sigma-1})^{\gamma} \sigma^{\frac{\sigma}{\sigma-1}}} \frac{\sigma-1}{\gamma-\sigma-1} \right]^{\frac{1}{\gamma}} \tag{39}$$

which is identical to V_j in section 6 in Redding (2011) and equation (48) in the Supplementary Appendix to Redding (2011). It follows then that the change in welfare from an international trade-policy liberalization (holding constant labor L_j, domestic fixed costs A_{jj}, and entry costs f_j^e) turns out to be:

$$d \ln V_j = (-1/\gamma) d \ln \lambda_{jj}. \tag{40}$$

which is identical to that in Arkolakis, Costinot, and Rodriguez-Clare (2012) and γ is the Pareto shape parameter in our model.

Having estimated the heterogeneous partial trade effects ($\widehat{\beta_{ij}}$) in section 3, we now start with actual bilateral trade flows and populations to generate general equilibrium effects on trade flows and wage rates (and hence national incomes) of EIA changes. We used actual bilateral flows from year 2005 for initial values of international flows X_{ij}. Since gross output is not available for aggregate goods production, as gross domestic product is a value-added measure of output, we needed to compute initial values of intranational trade using a gravity equation.[34] Armed with international and intranational trade flows, the market-clearing conditions are used to ensure prices and wage rates adjust to equate total trade with value-added output.[35] Using population and GDP data, initial wage rates are set equal to per capita GDPs.

Analogous to Head and Mayer (2014), we start with equation (17) from section 2, take the logarithm of both sides, and differentiate, replacing the last RHS term of (17) with the estimated partial effect ($\widehat{\beta_{ij}}$):

$$d \ln X_{ij} = d \ln w_j + \widehat{\beta_{ij}}. \tag{41}$$

For 2,266 bilateral EIA liberalizations (ij), the $\widehat{\beta_{ij}}$ are first calculated using the EIA coefficient estimates and interaction terms' coefficient estimates from Table 4, column 3 alongside the demeaned levels of the various trade-cost variables (Z_{ij}) described in section 3. For each ij bilateral liberalization, we remove the EIA, introducing the partial (direct) effect of the EIA on $\ln X_{ij}$, which is $\widehat{\beta_{ij}}$ in equation (41).[36] To derive the general equilibrium welfare effect, we need to compute the change in the N countries' gross domestic outputs and consequently wage rates w_j. We then recompute

[34] As addressed in Bergstrand, Larch, and Yotov (2015), there are basically two ways to handle the absence of direct measures of intranational trade. The first is the use of manufactures data because both international trade flows and gross output data are available; this was used in Bergstrand, Larch, and Yotov (2015). However, the shortcoming is a smaller number of countries for a much shorter time period. The second, applicable for our international aggregate trade flows for a long time panel, is to impute intranational trade flows using internal measures of distance.

[35] By our construction, there are no trade deficits or surpluses.

[36] Note that for the reciprocal EIA liberalizations we account for the direct effect on X_{ji}.

the matrix of all trade flows X_{ij} for all N^2 values of ij (including the N jj), and then new values of w_j, and Y_j. We iterate using a dampening factor until the changes in wage rates and prices are essentially zero. The change in welfare for country j for an EIA removed between i and j, $d \ln V_{ij}$, is $(-1/\gamma) d \ln \lambda_{jj}$, where following Head and Mayer (2014) we assume $\gamma = 5$. We conduct this process 2,266 times for 2,266 bilateral EIA removals. Finally, every one of the 2,266 simulations yielded unique values for the N national wage rates w_j, supporting section 2's theoretical conjecture of unique wage rates.

4.2 General Equilibrium Welfare vs. Partial Effects

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

$$d \ln V_j = d \ln w_j - d \ln P_j = d \ln w_j - \ln \sum_{k=1}^{N} \lambda_{kj} d \ln p_{kj}. \tag{42}$$

In our context, it is useful to rewrite equation (42) as:

$$d \ln V_{ij} = -\lambda_{ij} d \ln p_{ij} + d \ln w_j - \sum_{k \neq i}^{N} \lambda_{kj} d \ln p_{kj} \tag{43}$$

where $d \ln V_{ij}$ denotes the (log) change in country j's welfare from an EIA with country i.

Given an EIA can lower both variable and fixed trade costs, equation (43) can be written as:

$$d \ln V_{ij} = -(1/\gamma) \lambda_{ij} \widehat{\beta_{ij}} + (d \ln w_j - \sum_{k \neq i}^{N} \lambda_{kj} d \ln p_{kj}) \tag{44}$$

or:

$$d \ln V_{ij} = -(1/\gamma) \lambda_{ij} \widehat{\beta_{ij}} + \chi_{ij}. \tag{45}$$

Equations (44) and (45) decompose the welfare effect into the partial effect (the first RHS term) and general equilibrium effects (the second RHS term). 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 (as suggested by this section's introductory quote), and γ influences the effect on welfare with higher (absolute) γ diminishing the welfare gain.

We estimate equation (45) 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, χ_{ij}, is multiplicative and rewrite equation (45) as:

$$d \ln V_{ij} \quad = \quad -(1/\gamma) \lambda_{ij} \widehat{\beta_{ij}} \Phi_{ij} \tag{46}$$

Taking the logarithm of equation (46) yields a log-linear equation suitable for OLS:

$$\ln(d\ln V_{ij}) \quad = \quad \delta_0 + \delta_1 \ln(\lambda_{ij}) + \delta_2 \ln(\widehat{\beta_{ij}}) + \ln \Phi_{ij}. \tag{47}$$

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

Table 8 reports the results of estimating equation (47) under four alternative specifications. Specification (1) is equation (47), but constraining the coefficients δ_1 and δ_2 to be equal. Column (3) shows that the coefficient estimate for $\ln(\lambda_{ij}\widehat{\beta_{ij}})$ is positive and statistically significant. Moreover, the coefficient estimate of 0.92 is very close to the expected estimate of 1. Variation in $\ln(\lambda_{ij}\widehat{\beta_{ij}})$ explains 83 percent of the variation in $\ln(d\ln V_{ij})$.

In specification (2) in column (4), we allow the coefficient estimates for $\ln\lambda_{ij}$ and $\ln\widehat{\beta_{ij}}$ to be unconstrained. Column (4) shows that $\ln\lambda_{ij}$ and $\ln\widehat{\beta_{ij}}$ have positive and statistically significant effects on $\ln(d\ln V_{ij})$. Both variables explain 85 percent of the variation in $d\ln V_{ij}$. This suggests that the correlation between the welfare effect and the direct partial effect is extremely strong. Naturally, we would expect a correlation because the welfare effect is a function of the partial effect. Nevertheless, our result suggests that general equilibrium factors play a limited role empirically relative to the heterogeneous partial effects in influencing welfare. Finally, the coefficient estimate on $\ln\lambda_{ij}$ is economically very close to unity, as suggested by theory, though it is statistically different from unity (at the 1 percent significance level).

Specification (3) in column (5) adds an importer fixed effect to account for general equilibrium effects. The R^2 value rises from 0.85 to 0.91 with the inclusion of the importer fixed effect. Moreover, there is no material change in the estimated coefficients relative to specification (2). For completeness, specification (4) in column (6) includes an importer fixed effect and an exporter fixed effect. As for specification (3), the R^2 value rises from 0.91 to 0.94 with the inclusion of the importer and exporter fixed effects. Once again, there is no material change in the estimated coefficients relative to specification (2).

On net, the results suggest that welfare changes for importer j from an EIA with exporter i are well-approximated by partial effect estimates $\ln\lambda_{ij}\widehat{\beta_{ij}}$. 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 $\ln\widehat{\beta_{ij}}$ 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.

4.3 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 $\ln\widehat{\beta_{ij}}$ 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 1,431 country-pairs and predicted correctly 97 percent of the remaining 1,145 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:

$$y* = \varsigma_0 + x\varsigma + \epsilon \tag{48}$$

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:

$$P(EIA = 1) = P(y* > 0) = H(\varsigma_0 + x\varsigma) \tag{49}$$

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 7. 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. Table 9 presents the results of five alternative specifications. The number of observations (for year 2005) is limited to 2,266, as these are the number of estimates of $\ln \widehat{\beta_{ij}}$ from our earlier results.

The first specification, shown in column (3), regresses $\ln P(EIA_{ijt})$ on $\ln \lambda_{ij}$, $\ln \widehat{\beta_{ij}}$, and a constant. Two results are worth noting. First, both variables have the expected qualitative relationship with $\ln P(EIA_{ijt})$; as $\ln P(EIA_{ijt})$ serves as a proxy for $d \ln V_{ij}$, we cannot assign a specific expected quantitative value for the coefficients. Second, we note that the pseudo R^2 value is 53 percent.

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 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 (4) adds the log of bilateral distance to the regression. Column (4) shows that – although $\ln DIST_{ij}$ helps to explain $\ln P(EIA_{ijt})$ – $\ln \lambda_{ij}$ and $\ln \widehat{\beta_{ij}}$ still have significant explanatory power. Moreover, adding only distance increases the explanatory power from 53 percent to 85 percent. 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 7). Column (5) shows that – although all these observables are statistically significant in explaining $\ln P(EIA_{ijt})$ – $\ln \lambda_{ij}$ and $\ln \widehat{\beta_{ij}}$ still retain significant explanatory power. Moreover, the coefficient estimates for $\ln \lambda_{ij}$ and $\ln \widehat{\beta_{ij}}$ hardly change at all. The specification in column (6) adds an importer fixed effect. As shown in column (6), this has no material effect on the explanatory power of $\ln \lambda_{ij}$ and $\ln \widehat{\beta_{ij}}$. Finally, for completeness, the specification in column (7) adds both an importer and exporter fixed effect. Although the coefficient estimate for $\ln \lambda_{ij}$ becomes negative and marginally significant, the coefficient estimate for $\ln \widehat{\beta_{ij}}$ has no material change.

4.4 Ex Ante Analysis

Finally, although our study has like others focused on *ex post* analysis, our framework has implications for *ex ante* trade and welfare analysis. In discussing Arkolakis, Costinot, and Rodriguez-Clare (2012), or ACR, Melitz and Redding (2014) address the sensitivity of potential welfare-gain estimates (from liberalizations) owing to slight departures from the ACR restrictions. *Ex ante* trade-policy predictions are not generalizable due to the endogeneity of the partial trade-cost elasticity. Our framework helps to address this. We have shown that partial EIA effects are systematically related to observable proxies for variable and fixed bilateral export costs. Consequently, partial EIA effects of future EIAs can be predicted based upon these systematic relationships.

Moreover, our analysis shows also that general equilibrium welfare gains are *well approximated* by partial equilibrium welfare gains, where the latter are determined entirely by initial bilateral trade shares (which are readily observable) and EIA dummy coefficient estimates (which are readily estimated). Consequently, *ex ante* analysis of trade and welfare gains may be facilitated with our approach. It is clear that the approximation will be less accurate if the economic size of signatories is large; in these instances, the terms-of-trade effects will be more prevalent. Due to space constraints, we provide only one interesting computational anecdote. Using our partial equilibrium welfare approach, we computed the welfare gain to the United States (US) and to the European Union

(EU) of the proposed Trans-Atlantic Trade and Investment Partnership (TTIP) and compared the estimates to those from a standard computable general equilibrium analysis of the TTIP in Francois, Manchin, Norberg, Pindyuk, and Tornberger (2013). Our model's welfare estimates for the US (EU) of 0.09 percent were in line with estimates based upon Francois, Manchin, Norberg, Pindyuk, and Tornberger (2013), lying between the tariff only and less ambitious liberalization scenarios.

5 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 *ex post* and *ex ante* welfare calculations in the context of the new quantitative trade models.

However, 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, more work needs to be done and perhaps with an eye to earlier efforts to measure and analyze the effects of non-tariff measures addressed in Anderson, Bergstrand, Egger, and Francois (2008), which provided underlying methodology for Berden, Francois, Tamminen, Thelle, and Wymenga (2010) and Francois, Manchin, Norberg, Pindyuk, and Tornberger (2013). Such improved methodology can likely augment new quantitative trade model estimates of trade liberalizations examined in papers such as Caliendo, Feenstra, Romalis, and Taylor (2015), which omit policy fixed export cost changes likely associated with EIAs especially for developing countries.

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6 Tables

Table 1: Summary of Theoretical Effects of Trade-Cost Levels on the EM, IM, and Trade Elasticities

	Ad Valorem Tariff-Rate Elasticity Effects		
	Extensive Elasticity	Intensive Elasticity	Trade-Flow Elasticity
Lower fr_{ij}	+	+	+
Lower A_{ij}^N	+	0	+
Lower A_{ij}^P	+	0	+
	Policy Fixed Trade-Cost Elasticity Effects		
	Extensive Elasticity	Intensive Elasticity	Trade-Flow Elasticity
Lower fr_{ij}	0	0	0
Lower A_{ij}^N	+	0	+
Lower A_{ij}^P	-	0	-
	EIA Coefficient Effects		
	Extensive Elasticity	Intensive Elasticity	Trade-Flow Elasticity
Lower fr_{ij}	+	+	+
Lower A_{ij}^N	+	0	+
Lower A_{ij}^P	?	0	?

Notes: See text.

Table 2: Expected EIA Variables' Coefficient Signs

1	2	3	4
Variables	Expected Sign Extensive	Expected Sign Intensive	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.

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)	7,259	0.4	1.0
5 (Common Market)	5,516	0.3	0.8
6 (Economic Union)	2,619	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 x 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.

Table 4

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
EIA_t	+	0.16 ***	+	0.07 **	+	0.23 ***
		(4.08)		(2.31)		(6.94)
$EIA_t * \ln \text{DIST}$	−	-0.15 ***	−	-0.09 ***	−	-0.24 ***
		(-5.18)		(-3.83)		(-9.88)
$EIA_t * \text{ADJ}$	−	-0.21 ***	+	0.24 ***	?	0.03
		(-3.00)		(4.23)		(0.58)
$EIA_t * \text{LANG}$	+	0.17 ***	0	0.02	+	0.20 ***
		(2.92)		(0.47)		(3.82)
$EIA_t * \text{RELIG}$	+	0.16 ***	0	0.07	+	0.23 ***
		(2.75)		(1.39)		(4.51)
$EIA_t * \text{LEGAL}$	?	-0.14 ***	0	0.02	?	-0.12 ***
		(-2.92)		(0.60)		(-2.80)
$EIA_t * \text{COLONY}$	?	-0.35 ***	0	0.10	?	-0.25 **
		(-3.04)		(1.02)		(-2.52)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.824		0.821		0.912
N		70,173		70,173		70,173

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.

Table 5

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
EIA_t	+	0.01	+	0.20 ***	+	0.21 ***
		(0.20)		(6.03)		(6.03)
$EIA_t * \ln \text{DIST}$	−	-0.06 **	−	-0.14 ***	−	-0.20 ***
		(-2.46)		(-5.44)		(-7.69)
$EIA_t * \text{ADJ}$	−	-0.28 ***	+	0.21 ***	?	-0.07
		(-4.30)		(3.40)		(-1.14)
$EIA_t * \text{LANG}$	+	0.15 ***	0	0.05	+	0.21 ***
		(2.76)		(1.04)		(3.81)
$EIA_t * \text{RELIG}$	+	0.09	0	0.21 ***	+	0.29 ***
		(1.56)		(4.08)		(5.55)
$EIA_t * \text{LEGAL}$	?	-0.09 *	0	-0.05	?	-0.13 ***
		(-1.88)		(-1.14)		(-3.02)
$EIA_t * \text{COLONY}$	?	-0.53 ***	0	0.16	?	-0.37 ***
		(-4.54)		(1.49)		(-3.19)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.821		0.799		0.896
N		103,147		103,147		103,147

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.

Table 6

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
FTA_t	+	0.09 **	+	0.11 ***	+	0.20 ***
		(2.33)		(3.30)		(5.94)
$FTA_t * \ln \text{DIST}$	–	-0.17 ***	–	-0.04 *	–	-0.21 ***
		(-5.57)		(-1.76)		(-8.12)
$FTA_t * \text{ADJ}$	–	-0.19 **	+	0.30 ***	?	0.12 ***
		(-2.24)		(14.46)		(16.76)
$FTA_t * \text{LANG}$	+	0.15 **	0	0.04	+	0.19 ***
		(2.21)		(0.79)		(3.31)
$FTA_t * \text{RELIG}$	+	0.26 ***	0	-0.06	+	0.20 ***
		(3.88)		(-1.07)		(3.45)
$FTA_t * \text{LEGAL}$	?	-0.17 ***	0	0.07	?	-0.10 **
		(-3.13)		(1.55)		(-2.13)
$FTA_t * \text{COLONY}$	?	-0.31 **	0	0.10	?	-0.21 *
		(-2.41)		(0.94)		(-1.89)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.824		0.822		0.912
N		70,173		70,173		70,173

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.

Table 7

(1) Variables	(2) Expected Sign Extensive	(3) Extensive	(4) Expected Sign Intensive	(5) Intensive	(6) Expected Sign Trade	(7) Trade
CU_t	+	0.10	+	0.51 ***	+	0.61 ***
		(0.49)		(2.99)		(3.46)
$CU_t * \ln \text{DIST}$	–	-0.23 **	–	0.10	–	-0.13
		(-2.04)		(1.03)		(-1.36)
$CU_t * \text{ADJ}$	–	-0.20	+	0.28 **	?	0.08
		(-1.27)		(2.11)		(0.57)
$CU_t * \text{LANG}$	+	0.58 ***	0	0.06	+	0.64 ***
		(3.61)		(0.45)		(4.62)
$CU_t * \text{RELIG}$	+	0.25	0	0.03	+	0.27 *
		(1.34)		(0.18)		(1.72)
$CU_t * \text{LEGAL}$	?	0.05	0	0.04	?	0.09
		(0.32)		(0.32)		(0.69)
$CU_t * \text{COLONY}$	?	-1.03 **	0	-0.14	?	-1.17 ***
		(-2.51)		(-0.41)		(-3.31)
Fixed Effects:						
Exporter-Year		Yes		Yes		Yes
Importer-Year		Yes		Yes		Yes
Country-Pair		Yes		Yes		Yes
R^2		0.824		0.822		0.912
N		70,173		70,173		70,173

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.

Table 8: Determinants of (Logs of) Welfare Changes

(1) Variables	(2) Expected Coefficient Value	(3)	(4)	(5)	(6)
$\ln \lambda_{ij}\widehat{\beta_{ij}}$	1	0.92 ***			
		(105.02)			
$\ln \lambda_{ij}$	1		0.99 ***	1.00 ***	0.93 ***
			(104.49)	(119.66)	(68.47)
$\ln \widehat{\beta_{ij}}$	1		0.51 ***	0.58 ***	0.58 ***
			(18.62)	(22.41)	(21.22)
Constant	?	3.10 ***	3.28 ***	3.63 ***	0.45
		(48.99)	(53.69)	(5.68)	(0.49)
Fixed Effects:					
Importer		No	No	Yes	Yes
Exporter		No	No	No	Yes
R^2		0.830	0.847	0.908	0.939
N		2,266	2,266	2,266	2,266

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively.

Table 9: Determinants of (Logs of) Probabilities of EIAs

(1) Variables	(2) Expected Coefficient Sign	(3)	(4)	(5)	(6)	(7)
$\ln \lambda_{ij}$	+	0.24 ***	0.04 ***	0.12 ***	0.14 ***	-0.12 **
		(19.76)	(4.63)	15.91	22.27	-1.97
$\ln \widehat{\beta_{ij}}$	+	1.55 ***	0.52 ***	0.65 ***	0.62 ***	0.52 ***
		(43.77)	(20.52)	(16.66)	(19.57)	(16.65)
$\ln \mathrm{DIST_{ij}}$	−		-1.44 ***	-1.44 ***	-1.48 ***	-1.68 ***
			(-69.52)	(-54.01)	(-65.03)	(-24.14)
$\mathrm{ADJ_{ij}}$	?			-0.66 ***	-0.92	-0.64 ***
				(-12.92)	(-21.54)	(-7.83)
$\mathrm{LANG_{ij}}$	?			-0.17 ***	0.13 ***	0.37 ***
				(-4.60)	(3.48)	(7.90)
$\mathrm{RELIG_{ij}}$	?			-0.20 ***	-0.28 ***	-0.13 ***
				(-5.30)	(-8.32)	(-3.36)
$\mathrm{LEGAL_{ij}}$	?			-0.26 ***	-0.16 ***	-0.04 ***
				(-9.22)	(-6.50)	(-1.21)
$\mathrm{COLONY_{ij}}$	?			0.96 ***	0.66 ***	0.48 ***
				(8.53)	(7.41)	(5.54)
Constant		0.29 ***	9.03 ***	9.96 ***	11.26 ***	8.97
		(3.62)	(67.57)	(56.85)	(26.62)	(18.80)
Fixed Effects:						
Importer		No	No	No	Yes	Yes
Exporter		No	No	No	No	Yes
Pseudo R^2		0.533	0.847	0.884	0.938	0.951
N		2,266	2,266	2,266	2,266	2,266

Notes: *, **, and *** denote $p < 0.10$, $p < 0.05$, and $p < 0.01$, respectively.

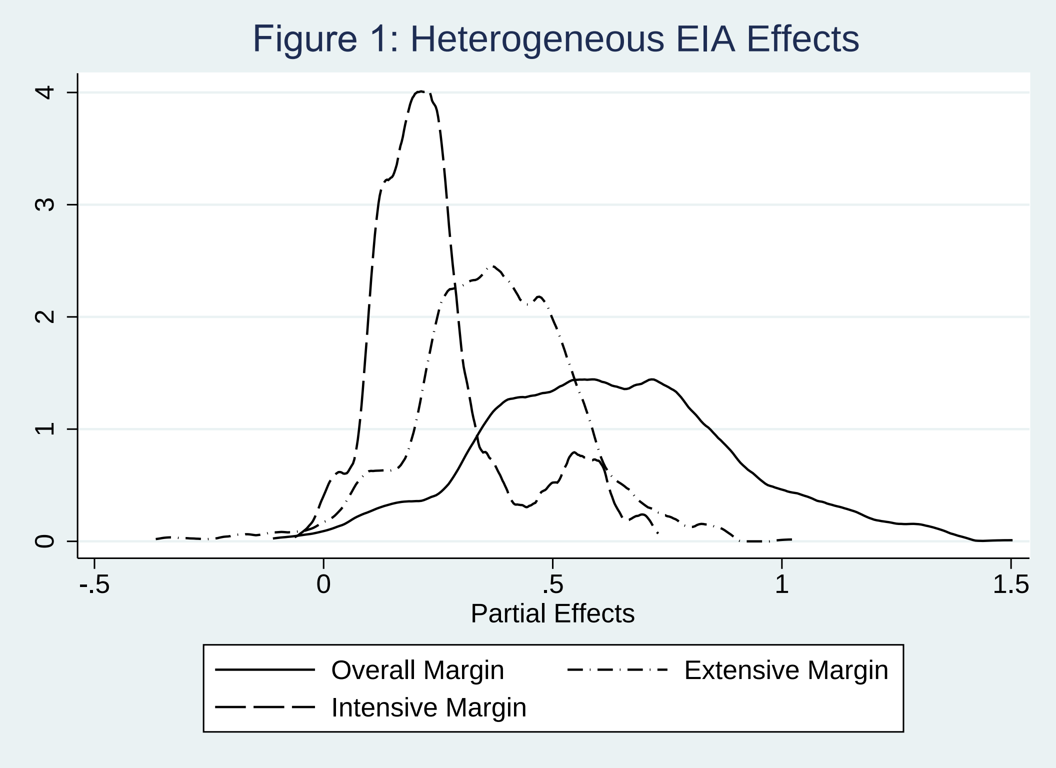
Figure 1: Heterogeneous EIA Effects
4
3
2
1
0
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1
1.5
Partial Effects
Overall Margin
Extensive Margin
Intensive Margin