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Modeling Heterogeneous Direct and Third-Country Effects of the Trade Policy Network

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ABSTRACT

This paper derives a structural Melitz-type gravity model featuring heterogeneous trade cost elasticities to estimate indirect and modular effects of trade agreements in a large panel dataset. Indirect or third-country effects are heterogeneous and nest trade diversion and reverse trade diversion. Our results show considerable heterogeneity of trade effects of agreements depending on pair-specific variable and fixed costs. For a single agreement, indirect partial effects are small relative to direct partial effects. However, a counterfactual simulation of the effects of the global network of agreements over 1972–2017 highlights that aggregate trade diversion can be substantial, because of the large and increasing number of existing agreements. Other factors like agreement depth, comparative advantage and specialization patterns of members also condition the potential for trade diversion of trade agreements.

JEL Classification: F13, F14, F15, F17, C23

1 | Introduction

The increasing number of international trade agreements, their varying content scope, and the diversity of members involved weave a complex web of trade linkages. It is well established that international trade liberalizations have positive aggregate trade, welfare and development effects.[1] Recent research explores how the direct effects of trade agreements vary across countries and agreements and which channels define this variation. It is less understood, however, how the indirect, third-country effects of the network of trade agreements vary and what economic implications they have.

This paper studies the heterogeneous indirect effects of Economic Integration Agreements (EIAs). It presents a theoretical framework to estimate heterogeneous modular effects from trade costs. Modular effects encompass direct and multilateral-resistance (third-country) effects and account for the trade impact of trade barriers holding output and expenditures constant (Anderson 2011; Head and Mayer 2014). Our framework expands the Melitz-type gravity model with firm heterogeneity and heterogeneous trade cost elasticities by Baier et al. (2018). We obtain the structural gravity model and derive approximated multilateral resistances, which provide a structural motivation for indirect or third-country effects—trade diversion

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and reverse trade diversion or external trade creation. Thus, indirect effects are based on pair-specific trade effects of EIAs. We augment the empirical model to include interactions of the EIA indicator and demographic variables, which, based on the urban economics literature, are proxies for distribution costs and effects of trade liberalizations through productivity. Moreover, we account for specific effects of the North American Free Trade Agreement (NAFTA) and the European Union (EU), which have a long history of economic integration and significant depth, and incorporate lagged effects that also enter the multilateral resistances (lagged indirect effects). We estimate the model for a large panel dataset and show a strategy to obtain unbiased estimates when using approximated multilateral resistances in panels.

Our estimates not only show relatively large heterogeneous bilateral effects but also that indirect effects of existing EIAs are heterogeneous, can be economically significant, and become more important over time. A counterfactual simulation of the effects of EIAs over our sample contributes to the understanding of the trade effects of EIAs over the last 45 years. Our simulation shows the quantitative importance of indirect EIA effects. On average, EIAs signed from 1970 to 2017 have a direct trade creation effect of 38%, while they divert trade by 11%, leading to a modular effect of about 18%. The simulation also provides insights regarding the nature of trade diversion, which depends on the levels of variable and fixed costs of the country pairs forming agreements. By isolating the effects of the NAFTA and the EU, and of the EU enlargement, we further show that other factors contribute to trade diversion effects such as the depth of the agreement, the number and size of members to an agreement, comparative advantage and specialization patterns.

This paper contributes to several branches of literature. First, it relates to recent research on heterogeneous effects of trade agreements. Novy (2013) documents substantial heterogeneity in trade costs broadly defined across country pairs. Baier et al. (2018) extend a standard Melitz model and show that the elasticity of trade flows associated with trade liberalizations is heterogeneous and endogenous to the levels of variable and fixed costs. They estimate pair-specific trade elasticities, captured by interactions between an indicator for EIA and geographic, institutional, and cultural gravity variables.[2] Given the structural motivation of the estimated heterogeneous elasticities, Baier et al. (2018) incorporate them in ex-ante simulations and show how accounting for heterogeneity of trade elasticities affects general-equilibrium inference about the welfare effects of trade liberalizations. Baier et al. (2019) emphasize the role of pairs' characteristics (within agreement variability), which they estimate produce 2/3 of the variation of the effects of Free Trade Agreements (FTAs), whereas the other 1/3 of the variation is attributed to variation across FTAs. They also conclude that FTA effects increase with ex-ante bilateral trade frictions (captured by pair fixed effects) and that the FTA effects on a country decrease with decreasing market power over terms of trade, with existence of previous trade agreements, and with increasing geographical distance. Our analysis builds on Baier et al. (2018) and expands the empirical model in two ways. First, we extend the model to provide insights on indirect effects—trade diversion and reverse trade diversion. Second, we include interactions capturing variable and fixed costs related to distribution and transport costs, specific scope and integration history of the EU and NAFTA, and lagged effects of EIAs.

Second, our research specifically relates to recent studies on the trade creation and diversion effects of trade agreements. These studies augment the gravity model with variables capturing third-country effects of membership to an agreement, but in most cases the introduction and definition of these variables lack a structural interpretation within the gravity model. Hence, the definition of the trade diversion variables used differs between studies.[3] In this context, a common problem is the collinearity between proxies for third-country effects and directional-time fixed effects capturing multilateral resistances such that the trade diversion variables often need to be excluded or re-defined to avoid collinearity or the directional-time fixed effects are excluded, potentially inducing bias.

Evidence from studies analyzing trade diversion is mixed. Carrere (2006) estimates some trade diversion effects on the importer side, whereas Magee (2008) finds limited evidence for trade diversion of average regional trade agreements, and Freund and Ornelas (2010) conclude that trade diversion is not a major concern for most regional agreements. By contrast, trade diversion is substantial on the importer side in Dai et al. (2014) and comparable to direct trade creation in Cheong et al. (2015). Also, trade diversion effects of agreements are smaller when a country is a member of preexisting agreements (Cheong et al. 2015, and Sorgho 2016) with marginal diversion effects decreasing in the number of agreements (Magee 2017). Furthermore, trade agreements can imply non-discriminatory liberalizations or harmonization of standards, which may have positive effects on third countries—namely, reverse trade diversion or external trade creation (Baldwin 2011). In this regard, Matoo et al. (2022) account for the depth of preferential trade agreements (PTAs) and distinguish between tariffs, discriminatory and non-discriminatory provisions. They find that non-discriminatory provisions have a reverse diversion effect, whereas tariff preferences and preferential provisions have a diversion effect. The diverting effects of tariffs depend on the depth of the agreement of the importer, i.e., deep agreements moderate the diverting effect of tariff preferences, which turns into external trade creation for deeper agreements.

We add to this research by providing a structural underpinning to both trade diversion and reverse trade diversion in terms of the components of the (approximated) multilateral resistances—i.e., our measures for trade diversion are derived from the multilateral resistances. Trade diversion emerges from the structural gravity model, and the multilateral resistances capture third-country effects of trade cost changes and absorb trade diversion effects of agreements. Moreover, third-country trade diversion and creation effects in our framework depend on variable and fixed costs, which are pair-specific and can be lagged.

Third, there is considerable evidence that trade liberalizations gradually affect trade flows.[4] Lagged direct effects of agreements can be rationalized by phase-in, price pass-through, and spatial effects of EIAs (Baier and Bergstrand 2007; Besedes et al. 2020). Our research adds to this evidence by identifying lagged bilateral and indirect effects of EIAs that are pair specific. We show that delayed effects of agreements also depend on the level of variable and fixed costs and are captured by lagged EIA interactions. More importantly, lagged heterogeneous indirect effects of EIAs appear in the multilateral resistances reflecting lagged trade diversion and reverse diversion effects. These lagged indirect effects can result from gradual phase-in of agreements, delayed price pass-through, and gradual market expansion of firms as a consequence of decreasing trade costs from agreements potentially combined with sequential exporting and learning (Albornoz et al. 2012, 2021; Morales et al. 2019).

Fourth, our derivation of third-country effects is based on a linear approximation of the multilateral resistances that is fully consistent with the structural definitions of output and expenditure (Fally 2015) and can be identified when using fixed effects in the estimation. In the context of cross sections, Baier and Bergstrand (2009, 2010) propose first-order Taylor approximations of the multilateral resistances which yield parameter estimates close to the unbiased fixed-effects estimator. Later research uses approximated multilateral resistances in panel data (Egger and Nelson 2011; Francois and Manchin 2013) and in gravity models with self-selection (Behar and Nelson 2014). Closer to our research, Cheong et al. (2015a) use approximated multilateral resistances and extract from them a trade diversion variable defined as the GDP-weighted sum of trade agreements signed by the importer. Egger and Pfaffermayr (2024) show that Baier and Bergstrand (2009)'s use of observed sales and expenditures is not consistent with the approximation point. This causes biased estimates, because the linear model does not approximate the nonlinear model at the approximation point. They propose a linear approximation based on a weighted projection matrix that corrects this bias. We show that first-order Taylor approximations using structurally-consistent (predicted) output and expenditure at the observed point (actual trade flows) are a specific case of the linear approximation proposed by Egger and Pfaffermayr (2024) and that our estimation approach using directional-time fixed effects is free from the bias induced by using incorrect approximations of the multilateral resistances in the estimation step. Moreover, Fally (2015) shows that when the data generating process is the structural gravity model, directional-time fixed effects in Poisson Pseudo-Maximum Likelihood (PPML, Santos Silva and Tenreyro 2006) identify the multilateral resistances. By contrast, we note that when drivers of trade flows other than those predicted by the theoretical model are present, directional-time fixed effects are required to obtain unbiased estimates in panels, and the approximation of the multilateral resistances identifies mathematically third-country effects conditioned on the structural model.

The paper is structured as follows. The next section presents the theoretical framework and the empirical model. Section 3 analyzes the econometric results. In Section 4, we study the modular, direct and indirect effects of EIAs over 1972–2017 based on a simulation. Section 5 concludes.

2 | Methodology

Our point of departure is the model for trade flows introduced by Baier et al. (2018), a Melitz-type model of trade with heterogeneous firms in spirit of Chaney (2008) and Redding (2011), which incorporates in an additive form policy and natural fixed costs, as well as endogenous fixed costs as in Krautheim (2012), and obtains pair-specific trade elasticities depending on the level of trade costs. We derive a structural gravity specification (see Anderson and Yotov 2010; Fally 2015, for a definition) from Baier et al. 2018's model and linearize the multilateral resistances. This model is estimated using the PPML estimator (Santos Silva and Tenreyro 2006) and allows for the estimation of the modular trade effect of trade barriers holding output and expenditures constant. The methodology is fully described in Appendix A.

2.1 | Derivation of the Structural Gravity

The model in Baier et al. (2018) can be summarized by three equations. The first equation characterizes trade flows from exporter i to importer j as

$$X_{ij} = \left[\left(\alpha_i L_i\right)\left(\varphi^*_{ij}\right)^{-\gamma}\right]\left[\left(\frac{\sigma\gamma}{\gamma-(\sigma-1)}\right) w_j \left(A_{ij} + M_{ij}^{-\eta}\right)\right] \quad (1)$$

L_i is the labor force (population) in country i, w_j is the wage in country j, and $A_{ij} = (A^P_{ij} + A^N_{ij})$ is the sum of policy and non-policy (exogenous) fixed costs. M_{ij} is the (equilibrium) mass of firms selling from i to j capturing endogenous fixed costs reflecting network effects (Krautheim 2012), and φ^*_{ij} is the productivity cutoff for firms selling from i to j. γ is the shape parameter governing the heterogeneity in the Pareto distribution of the firms' productivity, and $\sigma > 1$ is the elasticity of substitution of consumption across product varieties in the set of varieties Ω_j. α_i is an i-specific parameter depending on γ, σ and the exogenous fixed entry costs faced by firms in country i. The second equation is the zero-profit condition characterizing the equilibrium cutoff productivity level φ^*_{ij}:

$$\left(\frac{w_i \tau_{ij}}{\rho P_j}\right)^{1-\sigma} \frac{E_j}{\sigma}\left(\varphi^*_{ij}\right)^{\sigma-1} = w_j\left(A_{ij} + M_{ij}^{-\eta}\right) \quad (2)$$

where τ_{ij} are ad valorem iceberg variable trade costs ($\tau_{ij} \geq 1$), E_j is aggregate expenditure of country j, $\rho = (\sigma-1)/\sigma$ represents the firms' markup, $\eta \in (0,1)$ is the elasticity of fixed costs with respect to the mass of firms in i selling to j, and P_j, the equilibrium price index, is defined by the third equation:

$$P_j^{1-\sigma} = \sum_{i=1}^{N} M_{ij}\left(\varphi^*_{ij}\right)^{-\gamma}\left(\frac{w_i\tau_{ij}}{\rho}\right)^{1-\sigma}\left(\frac{\gamma}{\gamma-(\sigma-1)}\right)\left(\varphi^*_{ij}\right)^{\sigma-1} \quad (3)$$

Adding market clearing and trade balance conditions, the structural gravity model is

$$X_{ij} = \frac{Y_i E_j}{Y^w}\left[\frac{t_{ij}}{\theta_j \Pi_i}\right]^{-\gamma} \quad (4)$$

$$\theta_j^{-\gamma} = \sum_{i'=1}^{N}\left[\frac{s_{i'}}{\Pi_{i'}^{-\gamma}}\right] t_{i'j}^{-\gamma} \quad (5)$$

$$\Pi_i^{-\gamma} = \sum_{j'=1}^{N}\left[\frac{s_{j'}}{\theta_{j'}^{-\gamma}}\right] t_{ij'}^{-\gamma} \quad (6)$$

where Y_i and Y^w are output of exporter i and world output. θ_j and Π_i are the multilateral resistance terms, where i', j' refer to exporters and importers in the sample, and $s_{i'}$ and $s_{j'}$ are the weights of countries i' and j' in world output. t_{ij} is a catch-up variable capturing all trade costs, $t_{ij} = \tau_{ij}\left(F_{ij}\right)^{1/(\sigma-1)-1/\gamma}$, and F_{ij} collects fixed costs ($A_{ij} + M_{ij}^{-\eta}$). The system formed by (4–6) has the form of the traditional structural gravity model and is suitable for linearization of the multilateral resistances in spirit of Baier and Bergstrand (2009). In particular, we apply the linear Taylor-approximation using world shares based on predicted structural outputs and incomes. This produces expressions for the approximated multilateral resistances which are theory-consistent around observed trade costs.[5] That is:

$$\ln \theta_j = -\left[\sum_{i'=1}^{N} s_{i'} \ln t_{i'j} - \sum_{i'=1}^{N} s_{i'} \ln t_{i'1}\right] \tag{7}$$

$$\ln \Pi_i = -\left[\sum_{j'=1}^{N} s_{j'} \ln t_{ij'} + \sum_{i'=1}^{N} s_{i'} \ln t_{i'1} - \sum_{i'=1}^{N}\sum_{j'=1}^{N} s_{i'} s_{j'} \ln t_{i'j'}\right] \tag{8}$$

for $i, j = 2, \ldots, N$, with normalization $\theta_1 = 1$, and where $s_{i'}$ and $s_{j'}$ are theory-consistent weights of output of country i' and expenditure of country j', respectively, on world output and expenditure.[6]

The structural gravity Equation (4) can be cast in a form suitable for estimation using PPML

$$E\left[X_{ij}\right] = \exp\left\{\ln Y_i + \ln E_j - \ln Y^w - \gamma \ln \tilde{t}_{ij}\right\} \tag{9}$$

$$\ln \tilde{t}_{ij} = \ln t_{ij} - \left[\sum_{i'=1}^{N}\left(s_{i'} \ln t_{i'j}\right) + \sum_{j'=1}^{N}\left(s_{j'} \ln t_{ij'}\right) - \sum_{i'=1}^{N}\sum_{j'=1}^{N}\left(s_{i'} s_{j'} \ln t_{i'j'}\right)\right] \tag{10}$$

where $\tilde{t}_{ij}$ captures the modular effects of trade barriers on the pair ij, such that trade costs in t_{ij} enter both as bilateral trade costs, the first term in Equation (10), and as indirect (third-country) trade effects associated with the approximated multilateral resistances. These terms within the square brackets have a compelling mathematical form and constitute our measure of indirect effects (trade diversion and reverse trade diversion). The third term is the logarithm of the weighted geometric average of the trade costs across all pairs in the world, with weights equal to the product of the shares of the countries in the pair ij with respect to world output and income, respectively. The first and second terms within the square brackets are the weighted geometric average of the trade costs of all potential partners of the importer j and of the exporter i, respectively, weighted by the partners' share of world output and income, respectively.

These first two terms can also be expressed in terms of the weights associated with the pairs in the summations. The first (second) term is then the weighted geometric average of the trade costs of all potential partners of the importer j (exporter i), with weights equal to the product of the world shares of output and income of the countries in the pairs, and scaled by the ratio of world income (output) relative to the income of importer j (output of exporter i), a proxy for how large the potential market world demand (supply) is relative to the importer j (exporter i). Therefore, the modular trade effect of trade barriers between two partners depends on the bilateral trade barriers relative to the weighted geometric averages of trade barriers of each of the partners in the pair with their trading partners scaled by the size of the potential market faced by the importer and the exporter, respectively, and adjusted by a globalization trend captured by the (weighted geometric) average trade barriers worldwide including trade barriers between third countries.

2.2 | Empirical Specification and Estimation Strategy

For the empirical specification, it is convenient to substitute the variable t_{ij} capturing variable and fixed trade costs with a set of observables (see also Baier et al. 2018):

$$t_{ij} = \exp\left[\sum_{k=1}^{K} \beta_k z_{ij}^k EIA_{ij}\right] \tag{11}$$

EIA_{ij} is an indicator variable equal to one when there is an economic integration agreement in place between exporter i and importer j. z_{ij}^1 is a vector of ones, such that the first term in the sum only depends on the EIA_{ij}. z_{ij}^k $\forall \mathbf{k} > 1$ are observable variables, which represent geographical, cultural, institutional, historical and demographic factors related to variable and fixed trade costs. When such variable is an indicator variable, z_{ij}^k is equal to the variable,[7] while when it is not an indicator variable, $z_{ij}^k = Z_{ij}^k - \mu^k$ refers to the variable demeaned by its corresponding cross-sectional mean.[8] Equation (10) can be decomposed in a sum of analogous expressions for each of the components of t_{ij} in Equation (11). Therefore, the empirical model is:

$$X_{ijt} = \exp\left[\left(\sum_{l}\sum_{k} \delta_{kl} MTE_{ijl}^k\right) + \psi C_{ijt}\right] \exp\left\{\eta_{it} + \eta_{jt} + \eta_{ij}\right\} \exp\left\{u_{ijt}\right\} \tag{12}$$

where l is an index that allows to incorporate contemporaneous and lagged effects of EIAs $l = \{t-10, t-5, t\}$, and C_{ijt} is a vector of regressors including explanatory variables that are not absorbed by the fixed effects. The terms MTE_{ijl}^k refer to the modular trade effect of the interaction of EIA indicator with the variable z_{ijl}^k associated with period l, defined as

$$MTE_{ijl}^k = z_{ijl}^k EIA_{ijl} - \left[\sum_{i'=1}^{N}\left(s_{i't} z_{i'jl}^k EIA_{i'jl}\right) + \sum_{j'=1}^{N}\left(s_{j't} z_{ij'l}^k EIA_{ij'l}\right) - \sum_{i'=1}^{N}\sum_{j'=1}^{N}\left(s_{i't} s_{j't} z_{i'j'l}^k EIA_{i'j'l}\right)\right] \tag{13}$$

In Equation (12), the parameters associated with the terms capturing modular trade effects, MTE_{ijl}^k, are $\delta_{kl} = \gamma \beta^{kl}$, where $\beta^{1t} = 1$, such that $\delta_{1t} = \gamma$. η_{it}, η_{jt} and η_{ij} are exporter-time-, importer-time- and pair fixed effects, which capture unobserved heterogeneity, and u_{ijt} is the perturbation. The inclusion of it- and jt-fixed effects renders unbiased estimates robust to the selection of the country weights in the terms associated with third-country effects (see also Baier and Bergstrand 2009, 2010; Egger and Pfaffermayr 2024).[9]

Furthermore, the MTE_{ijl}^{k} variables in Equation (13) include indirect (third-country) effects, our measures for trade diversion, which vary on dimensions *it* and *jt* and thus are collinear with and absorbed by the *it*- and *jt*-fixed effects. Nevertheless, the parameter restrictions from the theoretical model—i.e., that the coefficients associated with third-country effects are equal to the EIA coefficients—identify the indirect effects through the bilateral EIA effect even when directional-time fixed effects are included. The identification of the effects of EIAs is facilitated by measuring them relative to intra-national trade (Yotov 2012; Borchert and Yotov 2017), which is not subject to bilateral EIA effects.

Our model incorporates lagged effects of agreements associated with specific variable and fixed costs and captured by the lagged EIA interactions. Lagged effects of agreements can be rationalized by phase-in, price pass-through, and spatial effects of EIAs (Baier and Bergstrand 2007; Besedes et al. 2020). The phase-in hypothesis (Baier and Bergstrand 2007) notes that EIAs remove a substantial part of trade costs gradually. Tariff phase-outs are typically implemented following stages defined in a schedule, and although some non-tariff barriers may be removed immediately as the agreement enters into force, other non-tariff barriers such as technical measures and standards may stay for longer periods until harmonization is achieved. Additionally, the delayed effect of the gradual phase-out of trade costs may be amplified when firms anticipate trade cost changes and smooth their purchases until trade costs are effectively reduced (Khan and Khederlarian 2021). The price pass-through hypothesis (Baier and Bergstrand 2007) postulates that tariff changes pass through to prices gradually over time, delaying the effects of the agreement on trade flows. The third explanation hinges upon potential spatial effects of EIAs (Besedes et al. 2020). As the implementation of the agreement removes trade barriers, exporting firms increase their profits and can face new costs associated with geographic expansion within the same region and gradually increase trade flows. Moreover, lagged heterogeneous effects of EIAs imply in our model lagged third-country, trade diversion and reverse diversion, effects appearing in the multilateral resistances. These lagged third-country effects are also triggered by gradual phase-in of agreements, delayed price pass-through, and gradual market expansion of firms as a consequence of decreasing trade costs from agreements, potentially combined with sequential exporting and learning (Albornoz et al. 2012, 2021; Morales et al. 2019).

2.3 | Observable Proxies for Variable and Fixed Trade Costs

The interactions of the observable variables with the EIA indicator introduce heterogeneity in the effects of EIAs at the pair level, such that the bilateral and third-country trade barriers—impliying direct and indirect trade effects—are heterogeneous pairwise depending on the proxies for variable and fixed trade costs. Following Baier et al. (2018), we include proxies for distance, adjacency, common language, religion similarity, common legal origin, and common colonial history, but we expand this set by including urban population share, urban population in major cities, and dummies for the EU and NAFTA.[10]

Distance (in logarithm) and adjacency are proxies for natural non-policy variable trade costs. Distance has an expected negative effect on both the extensive and intensive margins of trade and thus, an expected negative total effect. By contrast, the total effect of adjacency is a priori ambiguous. Adjacency has a positive effect on the intensive margin, as neighbor countries experience lower freight costs, but a negative effect on the extensive margin, when a negative relationship between border and the probability of trading is expected due to larger natural fixed costs (Helpman et al. 2008; Egger et al. 2011).

Common legal origin and common colonial history are proxies for institutional similarities. Institutional similarities are expected to have no impact on the intensive margin and an ambiguous effect on the extensive margin due to two forces (see Baier et al. 2018). On the one hand, a lower level of policy costs raises the share of endogenous export fixed costs in total exports fixed costs, which increases the variable-trade-cost and fixed-export-cost elasticities of EIAs. On the other hand, a lower level of institutional costs also lowers the relative importance of policy vs. non-policy fixed-export-costs, diminishing the policy fixed-export-costs elasticity of EIAs. That is, when two countries already have a common legal origin or a common colonial history, the potential of an EIA to reduce policy fixed-export-costs and thus the gains from liberalization are lower.

Religion and language capture bilateral cultural similarities, which are expected to decrease bilateral non-policy fixed trade costs and have a positive impact on the extensive margin and no impact on the intensive margin of trade (Baier et al. 2018). Accordingly, their total effects on bilateral trade are expected to be positive. However, the presence of a common language may relate to common legal and institutional origin. The introduction of a language in colonies is historically connected to the introduction of new values, rules and administration, and the new language serves as a way to establish the new rules and institutions and to facilitate their functioning. In this vein, Hall and Jones (1999) suggest that Western European countries helped themselves establish their institutional and legal structures, like property rights, in their colonies through the expansion of their influence and language. If countries share a common tongue as a result of their past colonial ties, language could reflect these institutional similarities. When language is related to legal and institutional links, its effect can reflect the effect of common legal origin and be negative. Therefore, in contrast to Baier et al. (2018), we regard the effect of common language as a priori ambiguous.

We also extend the model by including distribution and network effects associated with urban density. Firms benefit from urban density through increasing economies of scale and decreasing transportation costs, but there are also costs associated with urban agglomeration (see Duranton and Puga 2020, for a review survey). Urban density provides firms with the potential to realize economies of scale and to decrease costs associated with a more dense production and distribution network. Urban areas can reduce exporting costs at the origin and offer better distribution networks at the destination. This can be particularly relevant for firms' production within global supply chains. However, urban agglomeration also brings costs associated with higher rents and wages, which induce firms to locate outside urban areas, decreasing their potential to realize economies of scale and increasing costs associated with transport and distribution. These transport and distribution costs can be variable and fixed.

We incorporate interactions of the EIA indicator with two demographic variables associated with urban density to capture distribution and network effects. Urban population share captures decreasing costs from increasing urban density in a territory, and the share of population in major cities is used as a proxy for agglomeration costs.[11] Local trade costs applying to both imported and domestic goods—i.e., entering prices as a markup at the destination dimension—do not change relative prices to buyers and trade patterns in the gravity context (Anderson and van Wincoop 2004). Furthermore, exporting firms benefit from diminishing transport and distribution costs at both the origin and the destination and thus we define the proxies for urban density as the interaction between origin and the destination (including domestic sales). Since the variables related to urban density are defined at the pair level, they are not subject to the issue affecting local trade costs applying to both imported and domestic goods. Urban population and the proxy for agglomeration costs are also interacted with the EIA indicator, which is pair specific, to capture the international trade effect of EIAs depending on distribution and network effects associated with urban density. Additionally, these variables enter as controls without interaction and can affect domestic sales.

Moreover, urban density can correlate with the distribution of firms' productivity. Firms in urban areas may show larger productivity for several reasons (Duranton and Puga 2020). Increases in productivity translate into a larger mass of exporting firms, which expands the network effect emphasized in Krautheim (2012), in turn magnifying the effect of trade liberalizations.[12] Yet, excessive urban concentration has negative effects on productivity growth (Henderson 2003). This decreases the mass of exporting firms and reverts the network effect, lessening the effects of trade liberalizations. Because the relationships of our proxies for urban density to transportation and distribution costs and to firms' productivity have the same sign, the effects of our proxies can be predicted unambiguously. Thus, our proxies for urban density identify the trade effects of both distribution costs and productivity threshold changes associated with trade liberalizations. Urban population share and the agglomeration proxy are expected to increase and decrease, respectively, the effects associated with trade liberalizations.

Several mechanisms may channel endogeneity between trade and proxies related to urban concentration. Trade can affect the rate of growth of GDP, which in turn has effects on urban concentration (e.g. Williamson 1965; Ades and Glaeser 1995; Davis and Henderson 2003). Additionally, several factors link trade liberalization to urban density (see e.g. Monfort and Nicolini 2000; Bruelhart et al. 2004; Crozet and Koenig 2004), but other forces can attenuate this channel (Krugman and Livas Elizondo 1996; Behrens et al. 2007). The empirical literature reaches mixed conclusions concerning the relation between trade liberalization and urban concentration (see the survey by Bruelhart 2011 and Karayalcin and Yilmazkuday 2015). Yet, if trade levels can affect the process of trade liberalization, they can also affect urban concentration. The presence of directional-time fixed effects controls for the effect of trade on urban concentration through growth of GDP, while the inclusion of pair fixed effects controls for the endogeneity working through trade liberalization by accounting for invariant pair-specific factors that are the source of the endogeneity between trade and EIAs (Baier and Bergstrand 2007). The dynamics of these variables are mostly affected by history, long-term demographic patterns and urban policies and are exogenous to trade flows.

Finally, we include dummies for membership to the EU and NAFTA because of two main reasons. First, many country pairs belonging to these agreements are historically deeply integrated before the start of our sample or started integrating before the 1970s and in the first years of our sample (e.g., early agreements between some EU core countries and sectoral agreements between US–Canada). Second, the scope of these two agreements is large when compared with other agreements.

3 | Regression Results

We first present the estimation results of the gravity model with heterogeneous EIA effects. The regression analysis is based on several specifications run to achieve unbiased evidence and minimize efficiency issues from multicollinearity of the interaction terms.[13] Table 1 displays the main specifications. In columns 1 to 3, we compare specifications with approximated multilateral resistances with and without directional-time fixed effects, only including contemporaneous effects. In particular, we compare specifications with approximated multilateral resistances based on equal ($1/N$) weights and GDP weights (columns 1 and 2, respectively), including pair and time fixed effects, with a specification including directional-time fixed effects and pair fixed effects (column 3). The comparison shows that the correct estimation strategy when using approximated multilateral resistances includes directional-time fixed effects and uses structural parameter restrictions to identify the approximated multilateral resistances. Column 4 includes lagged variables but does not include dummies for the NAFTA and EU. Column 5 shows the benchmark specification, which adds lagged effects and directional-time fixed effects, and where insignificant variables are removed. The estimates confirm a positive baseline EIA effect and heterogeneity in contemporaneous and lagged EIA partial effects at the pair level depending on trade costs.

Focusing first on the estimation strategy, including directional-time fixed effects is required to control for unobserved heterogeneity at the country-time dimension despite the use of approximated multilateral resistances in the regression. While including directional-time fixed effects renders unbiased parameters' estimates in panels, previous research in the context of cross sections suggests that a specification including approximated multilateral resistances may not require the inclusion of country-time fixed effects. Baier and Bergstrand (2009) show supporting empirical evidence for this in a cross section, and Fally (2015) presents theoretical evidence suggesting that under PPML estimation the directional-time fixed effects map into output, expenditures and the multilateral resistances. The comparison of specifications including and excluding directional-time fixed effects (columns 1 and 2 vs. column 3 in Table 1) reveals that the coefficients associated with EIAs are biased under exclusion of directional-time fixed effects. The bias is large for the baseline EIA effect, which is four times larger under exclusion of directional-time fixed effects,

TABLE 1 | Trade gravity model: PPML main specifications.

	Approx. MRs (1/N)	Approx. MRs GDP	Three-way FEs	Three-way FEs and lags	Benchmark
$MTE_{ij,t}$	0.471***	−0.046*	0.132**	0.177***	0.186***
	(0.048)	(0.024)	(0.054)	(0.047)	(0.047)
$MTE_{ij,t}$ ln $DISTANCE_{ij}$	−0.058**	0.022	−0.082**	−0.043	−0.035
	(0.025)	(0.016)	(0.040)	(0.039)	(0.038)
$MTE_{ij,t}$ $ADJACENCY_{ij}$	−0.060	−0.250***	−0.210*	−0.227**	−0.280**
	(0.080)	(0.080)	(0.124)	(0.115)	(0.122)
$MTE_{ij,t}$ $LANGUAGE_{ij}$	−0.431***	−0.547***	−0.267***	−0.144**	−0.146**
	(0.068)	(0.069)	(0.079)	(0.070)	(0.069)
$MTE_{ij,t}$ $RELIGION_{ij}$	0.603***	0.192***	0.409***	0.327***	0.292***
	(0.076)	(0.072)	(0.100)	(0.105)	(0.101)
$MTE_{ij,t}$ $LEGAL_{ij}$	−0.009	0.198***	0.035		
	(0.055)	(0.050)	(0.064)		
$MTE_{ij,t}$ $COLONY_{ij}$	−0.144*	−0.276***	−0.031		
	(0.085)	(0.074)	(0.100)		
$MTE_{ij,t}$ $URBAN_{ij,t}$	−0.572***	0.959***	0.317*		
	(0.165)	(0.118)	(0.181)		
$MTE_{ij,t}$ $AGGLOMERATION_{ij,t}$	−0.368**	−0.057	−0.397**	−0.360**	−0.404**
	(0.144)	(0.101)	(0.158)	(0.158)	(0.159)
$MTE_{ij,t}$ $EU_{ij,t}$					0.195***
					(0.033)
$MTE_{ij,t}$ $NAFTA_{ij,t}$					0.329*
					(0.176)
$URBAN_{ij,t}$	5.267***	3.177***	−11.098***	−10.753***	−10.246***
	(0.214)	(0.238)	(1.943)	(1.947)	(1.795)
$AGGLOMERATION_{ij,t}$	2.099***	2.395***	−1.306***	−1.326***	−1.503***
	(0.195)	(0.186)	(0.453)	(0.445)	(0.441)
$MTE_{ij,t-5}$				0.039	0.026
				(0.027)	(0.027)
$MTE_{ij,t-5}$ ln $DISTANCE_{ij}$				−0.062**	−0.045*
				(0.024)	(0.024)
$MTE_{ij,t-5}$ $LANGUAGE_{ij}$				−0.173***	−0.154***
				(0.046)	(0.046)
$MTE_{ij,t-5}$ $RELIGION_{ij}$				0.180**	0.162**
				(0.080)	(0.078)
$MTE_{ij,t-5}$ $EU_{ij,t-5}$					0.085***
					(0.025)
Exporter-Year FE			Yes	Yes	Yes
Importer-Year FE			Yes	Yes	Yes
Country-Pair FE	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes			

Note: Standard errors in parentheses. *$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. Clustered standard errors at the pair level. The sample includes 326567 observations. The specifications with three-way fixed effects use Weidner and Zylkin 2021's bias correction. Approx. MRs stands for approximated multilateral resistances. In columns 1 and 2, the approximated multilateral resistances are based on equal weights ($1/N$ where N is the number of countries) and on GDP weights, respectively. Columns 3–5 include directional-time fixed effects, which allow for identification of parameters, and pair fixed effects (three-way fixed effects). Column 4 includes lagged variables but no interactions for EU and NAFTA. The benchmark specification (column 5) includes lagged variables and interactions for EU and NAFTA. Columns 4–5 exclude not significant terms. In columns 4 and 5, the contemporaneous distance interaction and lagged baseline EIA are not significant, but tests for joint significance of the contemporaneous and lagged effects in the benchmark specification do not reject jointly significant effects with p-values 0.027 and 0.0001 for distance and baseline EIA, respectively, and we include them both contemporaneously and lagged. Lagged terms beyond five years and forward EIA effects are not significant. See Table 7 in Appendix C for the results of alternative specifications.

and the point estimates of the interaction of urban population share change sign, while other interactions change significance.

One explanation for the bias when excluding directional-time fixed effects is that under misspecification of the structural model, directional-time fixed effects in PPML do not map perfectly into output, expenditures, and multilateral resistances. That is, when factors at the country-time dimension other than those in the theoretical framework drive international trade flows, the directional-time fixed effects account for their effects and are not fully explained by output, expenditures and the multilateral resistances. This is confirmed by the significant coefficients associated with other factors in a second-stage regression of the fixed effects on a set of variables, time and pair fixed effects (see Honoré and Kesina 2017), where we offset the multilateral resistances by taking advantage that the estimate of the first stage identifies the parameters associated with both the direct and third-country effects under the theoretical model (see Table 9 in Appendix C for results).[14]

Moreover, the inclusion of directional-time fixed effects eliminates concerns about the endogeneity of GDP appearing in the weights of the multilateral resistance terms and makes the estimation robust to the definition of GDP.[15] Accordingly, the correct empirical strategy includes directional-time fixed effects and takes advantage of the identification of the parameters associated with both direct and third-country effects of EIAs. Later calculation of the weights in the approximated multilateral resistances is based on predicted structural output and expenditure. Thus, under the structural restrictions to the parameters, the effects induced by multilateral resistances are fully identified and can be easily retrieved.

Turning to the estimation results, these show a positive baseline EIA effect and confirm the heterogeneity in EIA's partial effects. The baseline EIA effect, captured by the MTE coefficient, is positive and significant across specifications.[16] For the benchmark specification (column 5), the contemporaneous baseline effect of an EIA is 0.186 and significant, such that the baseline EIA partial effect increases trade flows contemporaneously by about 20.4%.

Although this effect is not directly comparable to average effects estimated in other research, it is slightly larger than the PPML estimates of the average EIA effect in Baier et al. (2018) and the contemporaneous coefficient of average FTAs estimated by Baier et al. (2019) in a model without pair-specific FTA effects. The lagged baseline EIA effect is 0.026 and increases considerably the point estimate of the baseline EIA.[17] The total baseline partial effect of an EIA becomes 23.6%. Our estimate is smaller than Baier et al. (2019)'s average effect, but a comparison is not straightforward, because Baier et al. (2019) do not include pair-specific effects in their specification and add dummies for globalization.[18] Also, a key difference is that Baier et al. (2019) estimate larger lagged than contemporaneous effects, whereas our estimates point to a larger contemporaneous effect.[19] However, our estimates also show that the EIA effects are to a large extent channeled through the pair-specific interactions.

The significant interactions support the existence of heterogeneous effects depending on trade cost levels and imply pair-specific channels that supplement the baseline EIA effect, increasing or decreasing the EIA effects at the pair level. Column 5 in Table 1 presents the interactions included in the benchmark specification. In general, the signs of the interaction effects match the theory.[20] Concerning the interactions with the two proxies for distribution costs, urban population share is not significant (and removed from the specification), whereas urban agglomeration is significant and decreases the effects of EIAs, consistent with the hypothesis that as the population is agglomerated in cities, the reduction in distribution costs is surpassed by increasing costs from agglomeration, such as high rent and wage costs, causing firms to locate in the outskirts of the cities and increasing distribution costs. The larger distribution costs imply smaller gains from joining an EIA.

The dummies for the NAFTA and the EU point to a larger effect of those agreements. Yet, their inclusion does not affect substantially the coefficients of the other terms (see columns 4 and 5 in Table 1). A substantial share of the effects of the EU are lagged, which may reflect specific features of the European common market that take longer to fully affect trade like competition policy and free movement of factors of production. The NAFTA seems to have a slightly larger effect than the EU. Yet, this may relate to the fact that the adjacency dummy has a negative sign, while the NAFTA implies preferential bilateral relations mainly between neighboring countries who have strong trade links.

For the rest of the interactions already present in Baier et al. (2018), the effects of EIAs decrease with distance and adjacency, and increase with common religion, which is consistent with Baier et al. 2018 Different to their findings, common legal origin and common colony, are not significant, and their exclusion in the model does not alter the other coefficients. Their effects seem to be captured by common language, consistent with the hypothesis that common language is associated with the existence of common legal and institutional origins, and which has a negative coefficient, suggesting that institutional similarities relate to lower policy fixed costs and tend to decrease the gains from trade liberalizations.[21] Moreover, there are significant pair-specific lagged effects of EIAs.[22] In the benchmark specification, the interactions with urban agglomeration, adjacency and the NAFTA enter only contemporaneously. The absence of lagged effects from urban population share and agglomeration suggests that urban agglomeration identifies variable and fixed costs related to distribution and transport costs but do not capture network effects associated with learning. By contrast, the interactions with distance, common language, religion and the EU enter both contemporaneously and lagged.

Overall, our estimates suggest several proxies for variable and trade costs which induce heterogeneity of the trade elasticities of EIA at a pair level consistent with Baier et al. (2018). They also point to an effect of common language in relation to common institutional frameworks and add a channel related to distribution costs captured by urban agglomeration and a different timing of the interaction effects reflecting the gradual implementation or lagged effects of EIAs.

To better illustrate the resulting heterogeneity in EIA effects, we analyze the bilateral partial effects of EIAs, which are defined as:

$$\frac{dX_{ijt}}{dEIA_{ijt}} = \exp\left\{\left(\gamma + \sum_k \sum_l \delta_k z_{ijl}^k\right)\left[1 - s_i - s_j + s_i s_j\right]\right\} - 1 \tag{14}$$

In Equation (14), the direct partial effect of an EIA is related to $\left(\gamma + \sum_k \sum_l \delta_k z_{ijl}^k\right)$, while the term $\left(\gamma + \sum_k \sum_l \delta_k z_{ijl}^k\right)\left[-s_i - s_j + s_i s_j\right]$ relates to the indirect effect of the EIA on the observed pair implementing it, which is channeled through the multilateral resistances. Trivially, the direct effects are larger than the indirect effects as long as $\left|-s_i - s_j + s_i s_j\right| < 1$.[23] Figure 1 displays the distributions of the direct, indirect, and modular partial effects associated with EIAs

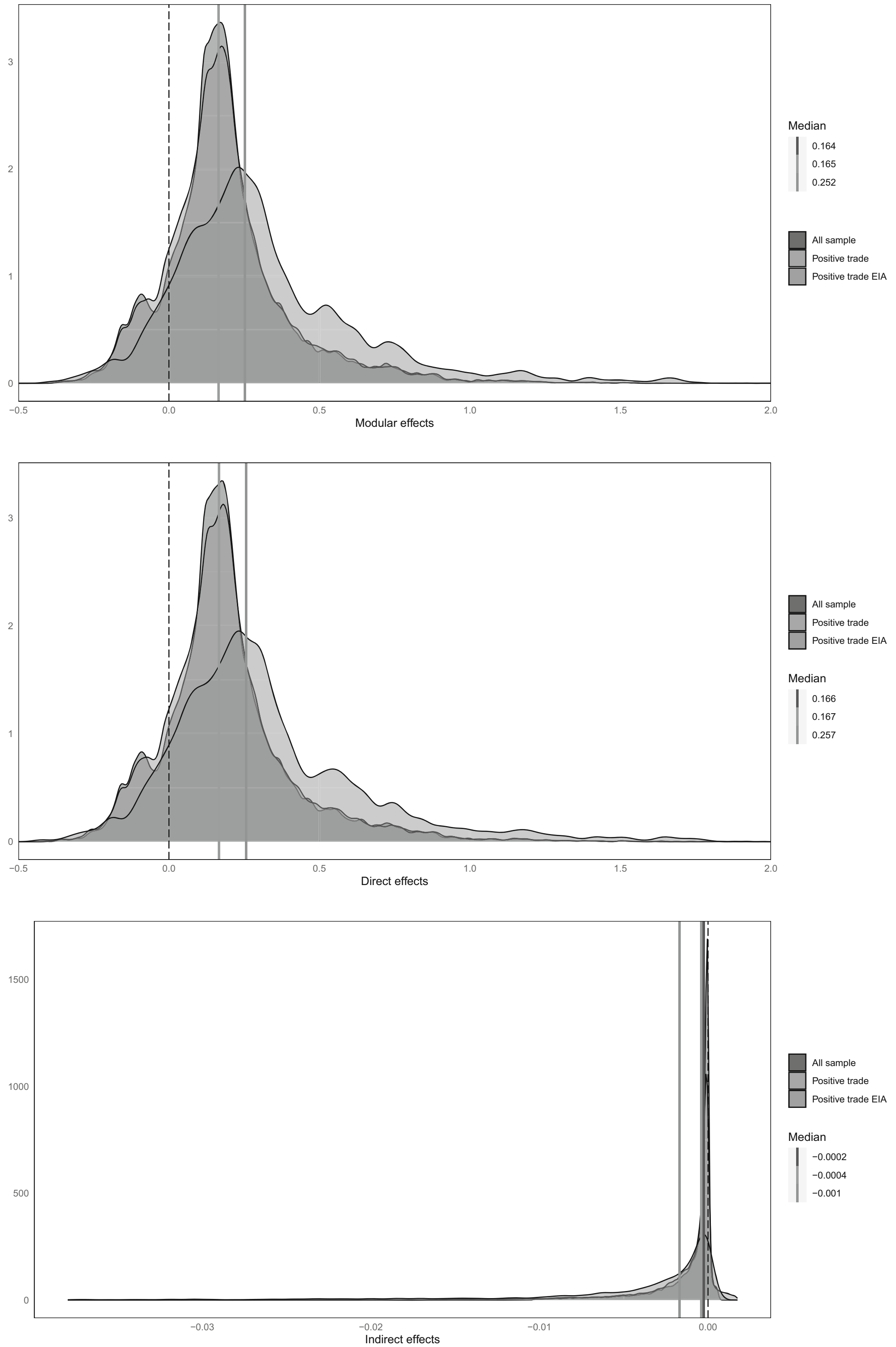

FIGURE 1 | Modular, direct, and indirect partial effects of EIAs. Effects calculated for a sample over 1972–2017. Modular, direct and indirect effects are shown in the figures on top, middle and at the bottom, respectively. The three densities in each figure reflect three distributions including different samples. All sample displays effects for all country pairs, positive trade shows the effects only for pairs with positive trade, while positive trade EIA shows the effects for pairs with positive trade flows and an existing EIA. [Colour figure can be viewed at wileyonlinelibrary.com]

at the pair level. The sample includes many zero trade flows and pairs without an EIA and thus, each plot contains three distributions, namely for all pairs (red), pairs with positive trade flows (green), and pairs with positive flows and EIA (blue).

The plots highlight the following findings. First, conventional wisdom attributes to EIAs a positive bilateral partial effect (trade creation) and a negative, small indirect effect (trade diversion), which we identify through the terms in the multilateral resistances. As we split the sample toward pairs with positive trade and existing EIAs, our estimates tend to agree with the conventional view on trade creation and diversion. The densities of the estimated pair-specific partial effects show that the share of observations conflicting with conventional wisdom is smaller as we change our focus from all observations to only considering observations with positive trade and an EIA in place. Second, the direct partial effects are rather heterogeneous and, although they are in general positive, 15% of the partial effects are negative when considering the whole set of observations. This percentage decreases to just 10% when only observations with positive trade flows and an EIA in place are considered. Third, the indirect partial effects are mostly negative—86% for the whole sample and 91% when only considering observations with positive trade flows and an EIA in place. Fourth, indirect partial effects are small relative to direct partial effects. Consequently, the distributions of the total partial effects are similar to those of the direct effects, and the shares of observations with negative total partial effects are the same as for the direct effects (15% and 10% when considering the whole set of observations and only observations with positive trade flows and an EIA in place, respectively). Therefore, the share of conflicting observations (negative direct effects and positive indirect effects) is relatively small when considering the support of the distributions. This highlights the importance of sample splitting to differentiate the relevant pair groups and to isolate EIA effects.

Our estimates also emphasize the importance of the lagged effects of EIAs. Figure 2 shows the distribution of total, contemporaneous and lagged partial effects of EIAs. The distribution of (5-year) lagged effects is more concentrated than that of contemporaneous effects, and although they are smaller, they explain a large share of the modular effects. Also, because of the positive correlation between contemporaneous and lagged effects, the distribution of total partial effects is more dispersed and presents a longer upper tail.

Notwithstanding the above, besides the indirect partial effects of a pair's EIA membership in Equation (14), as long as any of the members shares EIA partnerships with third countries, there emerge indirect effects on those EIA partnerships, as shown by the first and second terms in brackets in Equation (10). Moreover, the indirect effects also affect all trade partnerships worldwide, as the third term in brackets in Equation (10) suggests. This complex network of effects from multilateral agreements can be better represented through a simulation.

4 | Ex-Post Estimation of the Gains From Globalization

The intense process of international trade liberalization witnessed over the recent decades is characterized by two features: the proliferation of trade agreements and their multilateral character. Our model emphasizes the heterogeneous multilateral effects of EIAs and can be used to evaluate the direct, indirect, and modular trade effects of this process of trade liberalization from 1970 to 2017.

Prior research analyzes the general-equilibrium effects of globalization. Costinot and Rodríguez-Clare (2014) simulate the costs of autarky across countries and predict real income increases from trade liberalization between 1.5% and 8.1%, with an average of 4.4%. Anderson et al. (2018) simulate a border removal for all countries, obtaining exports increases ranging from 6.67% to 230%, with an average effect of 63%, and changes in real gross domestic product ranging from 0% to 10.31% with a moderate average effect of 2.63%. These experiments are defined in cross-sectional settings with samples of about 40 countries and a rest of the world aggregate.

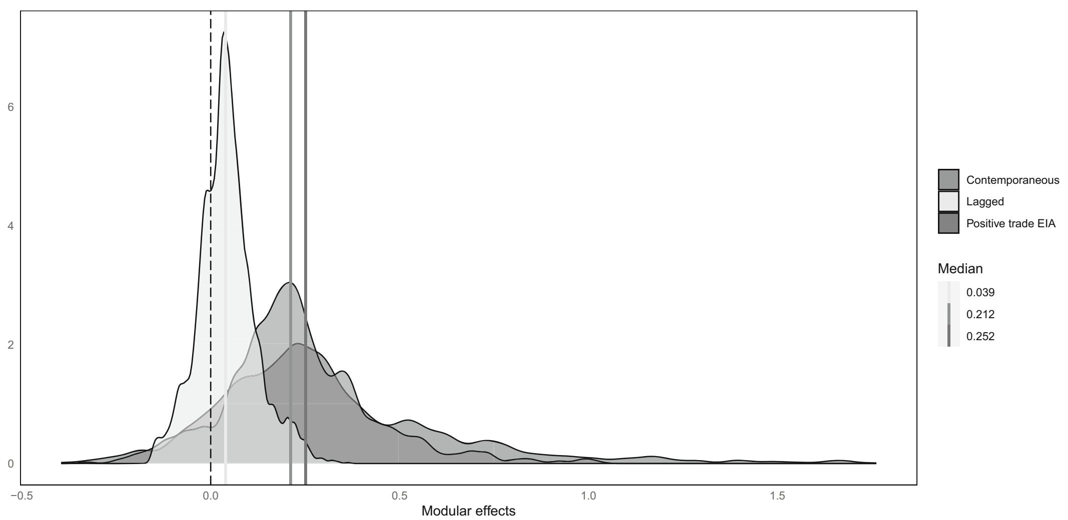

FIGURE 2 | Total, contemporaneous, and lagged partial effects of EIAs. Effects calculated for a sample containing pairs with positive trade and an existing EIA over 1972–2017. [Colour figure can be viewed at wileyonlinelibrary.com]

Our experiment extends previous research with respect to the number of countries (190 countries) and the period covered. It also expands previous settings by adding heterogeneous EIA effects conditioned on the regression estimates from Section 3, which include interactions capturing the dependence of EIA effects on pair-specific trade cost levels and lagged effects.

We estimate the long-term trade changes of joint memberships to EIAs in each period in the sample conditioned on the model parameters estimated and accumulating contemporaneous and 5-year lagged heterogeneous EIA effects. Our counterfactual exercise estimates the effects of entering into EIA partnerships across all pairs and each year in the period of analysis (1970–2017). As the model is static, the comparative statics reflect long-term trade changes. Moreover, our focus is the estimation of the indirect and direct trade effects of entering into EIA partnerships, and we leave out of the analysis price and income effects, as well as effects on factors of production. The estimated trade effects of EIAs contain both intensive and extensive margins. However, we do not model the probability of moving from zero to positive trade, such that this aspect of the extensive margin of trade is not accounted for and zero observations are not updated. In this sense, the estimates shown are conservative. The estimated effects of EIAs encompass tariff and non-tariff components of EIAs. We also provide a measure of the implied variation of trade frictions as a result of EIA formation. This requires an estimate of the tariff elasticity. We use the typical value of 4, which is also used by Baier et al. (2019) and allows for comparability.[24]

We first provide evidence on the relevance of indirect trade effects of EIAs relative to direct effects. After that, we analyze how other factors determine these direct and indirect trade effects. In particular, we address the role of economic development of the countries involved in EIA partnerships, the depth of an agreement and the number and size of the members of the agreement. To do so, we slice the sample of the estimated trade effects by level of development, by isolating the effects of the NAFTA and the EU, and by analyzing the effects of the enlargements of the EU from 2004 onwards for the different groups (members before 2004 and accession countries).

Table 2 displays the average trade effects for country pairs with a trade agreement across years and pairs, differentiating between direct, indirect, and modular effects. On average, EIAs signed from 1970 to 2017 have a direct trade creation effect of 38% while they divert trade by 11%, leading to a modular effect of about 18%.[25] These estimates are much larger than the 20% obtained by performing the same simulation using our sample but conditioning on the coefficients estimated by Baier et al. (2018) and only slightly larger than the average partial effect of 34% reported by Baier et al. (2019). Accordingly, assuming a tariff elasticity of 4, our model estimates an average 8.3% decline in trade barriers, which is very close to the 7% estimated by Baier et al. (2019), although the trade cost declines associated with modular effects are much smaller (about 4.2%).[26]

The model predicts positive direct bilateral effects of EIAs for 91% of the pairs and positive modular trade effects for 74% of the pairs. These estimates are larger than the 80% of positive direct trade effects calculated using our sample and the parameter estimates of Baier et al. (2018). These figures are also much larger than those reported by other research using heterogeneous effects models such as Baier et al. (2019), who estimate positive general-equilibrium effects for 57% of the sample, and Kohl (2014), who only find positive effects for 27% of the pairs.[27]

TABLE 2 | Summary statistics of trade effects of globalization.

Average effects accross years	Mean	Percentile	Equivalent trade friction
Direct	37.68	91	−8.32
Indirect	−11.45	45	2.99
Modular	18.39	74	−4.31
BBC direct	20.18	80	−4.70

Note: Effects calculated for a sample containing positive pairs with an existing EIA over 1972–2017. The number of pairs with an EIA is 14,053 out of a total sample of 326,567 observations. Direct, indirect and modular stand for direct, indirect and modular effects, respectively. BBC stands for results from the simulation conditioned on the coefficient estimates of Baier et al. (2018). The column percentile presents the proportion of positive effects. Equivalent trade frictions are calculated as $\left(1 - e^{-\overline{\phi}/4}\right) \times 100$, where $\overline{\phi}$ is the coefficient associated with the average trade effect in the first column. Average trade effects are $\left(e^{\overline{\phi}} - 1\right) \times 100$. Negative values correspond to declines in trade frictions.

Standard trade models with firm heterogeneity predict a positive trade effect of trade liberalization through two channels: market expansion and increased competition inducing inter-firm reallocation of production. However, the econometric model incorporating pair-specific effects predicts mostly positive but also some negative direct bilateral effects of trade liberalizations. Negative direct effects of EIAs may be related to sectoral reallocations and are also consistent with multilateral gains from joining a multilateral agreement that compensate for the negative effects of trade with a specific partner. These factors may explain the small share of negative direct trade effects (9%) found at the pair level. Also, the co-existence in the sample between deep and shallow agreements and between agreements that lack implementation (dormant) and those that have been effectively implemented may underlie some of the negative direct effects of EIAs estimated.

Indirect effects are on average about −11%, implying average trade diversion. Yet, there are positive, generally small, indirect effects for 45% of the pairs, highlighting that the network of trade agreements can generate third-country trade creation. In the standard model with firm heterogeneity, the interplay of market expansion and increased competition related to decreased trade costs associated with third countries provides an explanation for trade diversion. Yet, the pair-specific heterogeneity resulting from the interaction terms of our model allows for negative and positive third-country effects and provides evidence for the third-country trade diversion and creation effects noted by Magee (2008) and Baldwin (2011). Again, sectoral reallocations and the existence of shallow and deep agreements and of agreements that are dormant and those that have been implemented may underlie reverse trade diversion (external trade creation), but at least two other mechanisms can provide an explanation especially in the context of multilateral trade agreements, where multilateral (third-country) considerations are of particular interest when negotiating the agreement. The first mechanism refers to the reduction of non-discriminatory measures (Baldwin 2011; Matoo et al. 2022). If the trade cost reduction is associated

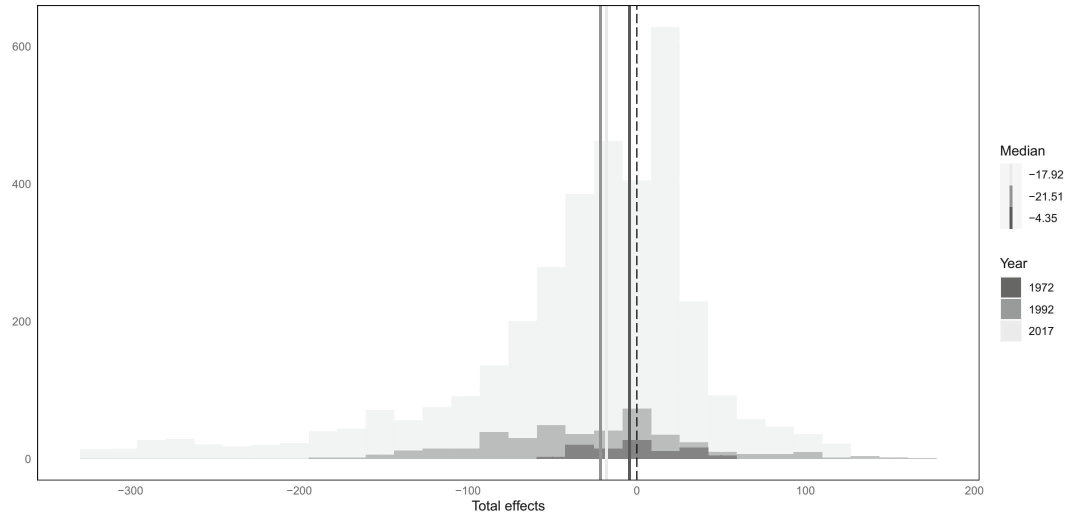

FIGURE 3 | Ratio of indirect to direct trade effects of globalization: Years 1972, 1997 and 2017. Ratios calculated for a sample containing positive pairs with an existing EIA. The histograms cover the central 94% of the distribution in each year to avoid outliers. [Colour figure can be viewed at wileyonlinelibrary.com]

with a non-discriminatory measure (e.g., non-tariff barriers like standards), third countries may benefit from these trade cost reductions, resulting in external trade creation. The second mechanism relates to the gains from the multi-destination character of exporting firms. Exporting firms tend to enter foreign markets that share similarities with their previous export destinations, because entry costs in new markets are smaller for firms that previously exported to similar countries (Morales et al. 2019).

A clear picture of the relevance of indirect effects over 1972–2017 is provided by the ratio of indirect to direct effects for country pairs with an agreement. This ratio shows the size of third-country effects that a country pair faces relative to the direct effects of their EIA partnership. Figure 3 presents the evolution of the distribution of the ratio for country pairs with an agreement for 1972, 1997, and 2017.[28] Two main findings emerge. First, the proliferation of EIAs strengthens the relevance of indirect effects relative to the direct effects of EIAs. Thus the distribution tends to move to the right for more recent periods, reflecting the increasing third-country effects as a result of the proliferation of multilateral EIAs in the last two decades. For a given country pair, the absolute value of the ratio increases as more EIAs become active and their third-country effects on the pair accumulate. The median of the distribution of the ratio changes from −5% in 1972 to almost −22% in 1992, a minimum of the sample, and to −18% in 2017, reflecting that indirect effects increase as more EIAs enter into force. Additionally, as more EIAs are active, the distribution of the ratio shows a larger lower tail. It also shows larger dispersion, because of the heterogeneity of the EIA effects resulting from the pair-specific interaction terms in the empirical model.

Second, most of the observations confirm a positive direct effect (trade creation), and the main split hinges on whether there is trade diversion or external trade creation. The ratio is negative when indirect and direct effects have different signs, being positive otherwise. In 2017, 54% of the observations show positive direct and negative indirect effects (trade diversion), as predicted by standard theory. For the whole sample, this share is maintained across all years and 19% of the observations present larger trade diversion than trade creation (larger negative indirect effects than positive direct effects). These cases seem related to high income countries which tend to enter into a significant number of EIA partnerships. This is the case of some EU countries and few EU neighbors. In 2017, 36% of the observations present a positive direct effect and positive indirect effects (external trade creation). Across all years, about 38% of the pairs with an EIA show positive direct and indirect effects. Countries that tend to have more occurrences within this group are small countries in the Caribbean and in the EU. In 2017, 6% of the observations show negative direct and positive indirect effects; and 2% exhibit direct and indirect negative effects.

The theoretical model predicts larger direct and indirect trade effects of EIAs, the lower natural costs and the larger policy costs. It is of interest to understand to what extent other factors determine direct and, more relevant to our purpose, indirect EIA effects. Thus, we turn to the relation between economic development and the effects of trade agreements in the context of our model. Conventional wisdom holds that policy costs are larger between developing economies, such that larger trade effects of EIAs are expected between pairs formed by developing countries. Prior research concurs with that conclusion. Anderson et al. (2018) suggest that a complete removal of international borders benefits large developed countries less, while Baier et al. (2018) conclude that less developed economies are likely to benefit from an agreement much more than developed economies. We slice the results of our simulation across the level of development of the pair's members (see Table 3). Trade effects of EIAs are conditional on pairs' level of economic development. On average, high-income country pairs benefit more through the direct effects of trade agreements and experience larger trade diversion, whereas low-income country pairs benefit less from direct economic gains but receive positive effects from third countries (reverse trade diversion or external trade creation). To the extent that economic development is positively related to the formation of agreements (see e.g. Baier and Bergstrand 2004), it is expected that pairs formed by high development countries have more EIA connections and experience the largest trade diversion.

TABLE 3 | Summary statistics of trade effects of globalization: Split by level of development.

	Globalization		High-High		Low-Low		High-Low		Low-High	
Average effects	**Mean**	**Percentile**	**Mean**	**Percentile**	**Mean**	**Percentile**	**Mean**	**Percentile**	**Mean**	**Percentile**
Direct	37.68	91	44.75	93	10.32	70	19.40	88	19.98	88
Indirect	−11.45	45	−15.61	33	7.31	99	−1.89	70	−2.02	69
Modular	18.39	74	18.70	72	18.32	80	16.90	81	17.38	82

Note: Effects calculated for a sample containing positive pairs with an existing EIA over 1972–2017. Direct, indirect and modular stand for direct, indirect and modular effects, respectively. The column percentile presents the proportion of positive effects. The distinction between high and low income countries is based on the classification by the World Bank (based on GDP per capita). High-High: Pairs in which both countries belong to high or upper middle income. Low-Low: Pairs in which both countries belong to low or lower middle income countries. High-Low: Pairs where the exporter is high or upper middle income country and the importer is a low or lower middle income one. Low-High: Pairs where the exporter is a low or lower middle income country and the importer is a high or upper middle income country.

Still, the positive relationship between development and the size of the direct trade effects of EIAs is in sharp contrast to prior research. Moreover, this relationship underlies the formation of the measures of the indirect trade effects. Therefore, we further investigate this relationship.

We regress, alternatively, the (logged) pairwise predicted direct and indirect trade effects on GDP, and on GDP per capita (GDPpc) and population (the results are shown in Table 13 in Appendix D). The regressions for indirect effects confirm that trade diversion increases with GDPpc and country size. With respect to the direct effects, GDPpc captures the level of development while population refers to the size or scale of a country. The development effect is expected to be positive. The more developed countries, being more competitive and showing larger potential consumption per capita, experience larger benefits after a reduction of trade barriers. In this vein, Francois and Manchin (2013) point out that higher infrastructure and institutional quality are positively correlated with development levels and trade flows. The size effect is expected to be negative, such that larger countries, as captured by population, experience lower increases in trade after a reduction of trade barriers. The correlation analysis suggests that EIAs' trade direct effects are positively correlated with GDP and GDPpc and negatively correlated with population (see Table 13 in Appendix D). In this regard, Anderson et al. (2018)'s negative correlation between GDP and trade effects of a border removal, which is estimated using a cross-section of 40 countries, may capture the mix of the two different effects. However, Baier et al. (2018) report a negative correlation between predicted trade effects and GDPpc significant at the 10% level. The difference with our results may be related to the specific treatment of the EU and NAFTA in our model.[29] These agreements are deeper and more extensive and are expected to be more effective in promoting trade (Baier et al. 2014). To the extent that this is the case and the depth of trade agreements relates to economic development (see Kohl et al. 2016), the fact that we account specifically for the EU and NAFTA may underlie the positive relation found between economic development and predicted direct trade effects.

Therefore, we zoom in on the effects of the EU and NAFTA in Table 4, which shows the direct and indirect effects experienced by their members. Both agreements show larger direct effects than the rest of EIAs and substantial trade diversion for countries within each agreement, which renders modular effects that are much smaller than the direct effects. With regard to the NAFTA, the direct effects are sizable and positive for all its members. The average direct effect is slightly smaller than for high income countries (shown in Table 3), which includes the effects of the EU, but higher than the effects of other agreements (not shown), and in line with the estimates of Caliendo and Parro (2015). However, given the large trade diversion generated, in part induced by the large size of the partners forming the agreement, the average modular effect is about two thirds of the direct effects. Turning to the EU, it generates much larger direct effects, which are on average about 91% and very heterogeneous, with some pairs showing effects above 186%. The larger effects can be explained by the ability of our model to capture the deep and extensive character of the EU through pair-specific heterogeneity of the EIA effects. Yet, the large number of members and the potential for large trade diversion, also associated with the depth of the agreement, downsize the effects significantly and yield modular effects around 25%, which are smaller than the average (general-equilibrium) trade effects reported by Mayer et al. (2019) and Felbermayr et al. (2022), which are around 100% and 46%, respectively.

TABLE 4 | Summary statistics of the EU and the NAFTA.

	NAFTA		EU	
Average effects	**Mean**	**Percentile**	**Mean**	**Percentile**
Direct	32.27	99	91.15	99
Indirect	−9.56	—	−34.09	8
Modular	20.33	68	25.94	64

We further analyze the effects of the EU enlargement from 2004 onwards and slice the effects of the EU in four groups depending on whether countries were part of the EU before 2004 (EU15) or part of the successive enlargements taking place from 2004 onwards (EUacc).[30] Given the differences in economic development between the EU15 and the new EU members, one can learn about the relationship between the trade effects of agreements and economic development of their members, controlling for the content of the agreement and the number and size of its members.

Table 5 summarizes the effects. The enlargement entails large gains for the pairs involved. The direct effects are slightly larger than for the average EU effect in Table 4, and the trade diversion generated is relatively smaller, such that the average modular effects generated by the enlargement are considerable—i.e., around 50% for pairs consisting of EU15 and new EU members and above 60% for pairs formed by new EU members. These effects are larger than the effects found by Egger and

TABLE 5 | Summary statistics of the EU enlargements from 2004 onwards.

	Enlargement		EUacc-EUacc		EUacc-EU15		EU15-EU15	
Average effects	**Mean**	**Percentile**	**Mean**	**Percentile**	**Mean**	**Percentile**	**Mean**	**Percentile**
Direct	95.27	99	91.19	99	96.81	99	—	—
Indirect	−21.18	0	−13.61	1	−24.04	0	−33.45	—
Modular	53.34	93	64.52	99	49.12	91	−33.45	—

Larch (2011), who show increases among new EU members around 30%, and by Wolfmayr et al. (2019), who report 48% and 6.6% increases for the new EU members and the EU15, respectively. Yet, assuming a tariff elasticity of 4, the estimated effects translate into declines of trade barriers of circa 20% and 13% for direct and modular effects, respectively, consistent with the range of trade reductions considered by Baldwin et al. (1997). The EU enlargement also generates substantial trade diversion for the EU15 countries and implies sizable net trade diversion from the EU15 to the enlargement countries (see also Egger and Larch 2011). These trade creation and reallocation reflect the increased integration of the new EU members in European supply chains as a result of specialization patterns and lower production costs and factor prices.

Furthermore, the relationship between the trade effects of the agreement and the level of economic development of its members shows a complex pattern. Direct effects are larger for pairs formed by EU15 and new EU members, while pairs consisting of two accession countries show relatively lower direct effects. Using the average direct effect of the EU15 in the sample year before the enlargement (2002) as a proxy for trade effects for pairs formed by high development countries, the trade creation effect is slightly smaller (about 90%). Nevertheless, the different trade diversion generated across groups in Table 5, much smaller for pairs consisting of two accession countries, results in larger modular effects for pairs formed by new EU members. All in all, this indicates that additional factors like comparative advantage affect the relationship between economic development and the trade creation and diversion effects of trade agreements, and that the scope for substitution of production from third countries determines the potential for trade diversion of an agreement.

5 | Conclusions

The proliferation of trade agreements over the past decades and their increasing multilateral character define the complex world trading system and determine global interconnectedness of countries. Our work aims at understanding the nature and empirical magnitude of indirect, third-country effects of preferential trade agreements in the context of globalization. These indirect, trade diversion effects of trade agreements may underlie PTA-contagion and the spread of regionalism, as one of the motivations for countries to sign PTAs is to reduce the discrimination effects of trade agreements signed by their partners (see Baldwin 1993; Baldwin and Jaimovich 2012; Egger and Larch 2008).

This paper derives structural measures of the indirect EIA effects in the framework of a Melitz model (Baier et al. 2018) that predicts pair-specific trade elasticities of EIAs depending on the level of variable and fixed costs. Indirect EIA effects are captured by approximated multilateral resistances which aggregate pair-specific EIA effects and encompass both trade diversion and external trade creation. The structural gravity model is estimated for a panel dataset covering 190 countries between 1972 – 2017 and allows to estimate direct, indirect and modular effects of EIAs.

The regression results provide evidence for heterogeneous EIA effects depending on pair-level variable and fixed costs. The results point to channels not identified by previous research, related to urban density, to common language as a proxy for common legal and institutional systems, and to the historical integration and larger depth of the EU and NAFTA. The partial effects confirm the heterogeneity of EIA effects and the importance of lagged EIA effects. For a single EIA, the indirect partial effects, as second-round effects, are small relative to the direct partial effects. Nevertheless, a simulation of the effects of the global network of EIAs highlights that indirect trade effects experienced by country pairs can be substantial, because of the large and increasing number of existing agreements as well as the large number of countries involved in preferential partnerships.

On average, aggregate trade diversion halves the direct effects of EIAs and produce substantially smaller modular effects. The average size of indirect relative to direct effects increases over the sample and is about 18% in 2017. As the econometric model shows, trade diversion depends on the levels of variable and fixed costs of the country pairs forming agreements. Motivated by the relationship between direct and indirect trade effects and economic development, we analyze the trade effects of EIAs isolating depth of the agreement, and number and size of members. This analysis suggests that trade diversion strongly relates to the depth of agreements, and to the number and size of the partners with further agreements. Moreover, our findings indicate that trade diversion is associated with comparative advantage and specialization patterns, and with the potential for substitution of goods across origins. To the extent that these factors drive the potential for trade diversion of trade agreements, they will underlie the decision to sign new trade agreements against intrinsic resistance to membership and thus, will drive how regionalism spreads. Future research should factor in these potential drivers formally.

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Conflicts of Interest

The authors declare no conflicts of interest.

Data Availability Statement

The data that support the findings of this study are available from the corresponding author upon reasonable request.

Endnotes

[1] See e.g., Anderson and Van Wincoop (2003); Eaton and Kortum (2002); Melitz (2003) and Costinot and Rodríguez-Clare (2014).

[2] Baier et al. (2018) provide empirical support for heterogeneous trade elasticities at the extensive and the intensive margins, as well as of total trade. Our research focuses on modeling third-country effects without distinguishing between third-country effects at the extensive or intensive margins. The extension of our framework to determine heterogeneous third-country effects at the extensive and intensive margins is straightforward.

[3] The analysis of trade creation and diversion dates back to the seminal articles by Viner (1950) and fundamentally to Viner (1950) and includes a large number of references. We focus on the most recent research, mainly in the framework of panel data. Carrere (2006) uses agreement-specific dummies representing trade diversion from the importer and exporter sides. Magee (2008) uses the count of the number of agreements of observed trading partners with third countries for a generic trade agreement and specific agreements. Cheong et al. (2015) modify Magee (2008)'s count to avoid multicollinearity with the directional-time fixed effects. Cheong et al. (2015) define trade diversion in terms of the GDP-weighted sum of trade agreements signed by the importer while controlling for approximated multilateral resistances. Dai et al. (2014) distinguish between diversion affecting importer and exporter sides. Matoo et al. (2022) use the relative preference margin and the average depth of trade agreements on the importer side, and both measures are defined such that they can be included together with the importer-time fixed effect.

[4] See e.g., Baier and Bergstrand (2007); Olivero and Yotov (2012); Baier et al. (2019); Besedes et al. (2020); Khan and Khederlarian (2021).

[5] Egger and Pfaffermayr (2024) show the link between the approximation in Baier and Bergstrand (2009) and a linear approximation based on a weighted projection matrix. We are thankful to Peter Egger for sharing their paper. In Appendix A we show that this linearization produces expressions for the multilateral resistances analogous to ours when the components of the weighting matrix are equal to the weights we use.

[6] To compute the theory-consistent weights, output and expenditure follow a structural definition, such that for the exporter and the importer they are calculated as the sum of predicted trade flows across trade partners (see Anderson and Yotov 2010, and Fally 2015) and thus, they are theoretically and econometrically consistent with the structural model estimated. See Appendix A for further details.

[7] This is the case in our empirical analysis for adjacency, common language, common legal origins, common colony, EU and NAFTA.

[8] In our empirical analysis, these variables are distance, common religion, urban share, and agglomeration.

[9] That is, the definition of the weights in the MTE_{ijl}^{k} variables does not affect the coefficients' estimates as the third-country effects are absorbed by the importer-time and exporter-time fixed effects.

[10] For details on the variables' definition and sources see Appendix B.

[11] The inclusion of pair fixed effects in our specifications brings the reading of our proxies closer to urban density, the theoretical counterpart.

[12] Specifically, the increasing network effect resulting from the larger mass of exporting firms magnifies the ad valorem tariff-rate elasticity of the extensive margin but does not make the elasticity endogenous (Krautheim 2012).

[13] This is relevant to optimize the efficiency of the model's prediction in the simulation of counterfactuals.

[14] In the second-stage specifications, several variables are significant, confirming that the directional-time fixed effects are not fully explained by output, expenditures and multilateral resistances. Plenty of empirical evidence supports the inclusion of other determinants running on the dimensions of the directional-time fixed effects in ad-hoc gravity regressions. We restrict our estimations to a set of variables sufficiently large to avoid large misspecification issues while keeping broad country coverage, which is of relevance for the simulation in Section 4.

[15] Baier and Bergstrand (2009, 2010) and Egger and Nelson (2011) suggest using the inverse of the number of countries ($1/N$) to avoid endogeneity between GDP-based shares and trade flows. Anderson and Yotov (2010) and Fally (2015) favor using structural output and expenditure based on the sum of trade flows to keep consistency with the theoretical model. Egger and Pfaffermayr (2024) show that the selection of weights can add bias to the approximation bias. As the fixed effects in our specification absorb the multilateral resistances, the parameter estimates are robust to (do not depend on) the weights' definition. Thus, it is unnecessary to use uniform or GDP-based weights, and the elasticities associated with the multilateral resistances are fully identified by the direct effects of EIAs under theoretical restrictions.

[16] The baseline EIA effect corresponds to an EIA for which the interactions using indicator variables are evaluated at zero and the interactions for continuous variables are evaluated at the cross-sectional mean. Thus, it is not an average effect.

[17] Although the lagged baseline EIA effect is not significant, it is jointly significant with the contemporaneous effect and improves the model estimate. Removing it increases the contemporaneous parameter by a similar magnitude but decreases the model's likelihood.

[18] To test for robustness of the results and in spirit of Bergstrand et al. (2015), we also run estimations that include international trade border dummies yearly to isolate whether there is a globalization trend. Our results are robust to their inclusion and these dummies are insignificant (see Table 11 in Appendix C).

[19] In general, PPML panel estimates tend to be smaller than OLS panel estimates (see also Anderson and Yotov 2010), and the baseline EIA effects found are smaller than OLS estimates using panel data in (Baier et al. 2018; Baier and Bergstrand 2007; Bergstrand et al. 2015) but close to Novy (2013).

[20] Table 7 in Appendix C shows a full-set specification including all the interactions. The results are qualitatively similar to those from the benchmark specification, although, due to multicollinearity, some of the interactions are not significant.

[21] The correlation between common language and common legal origin and common colony is of 0.5. It should be noted that Baier et al. 2018's estimates using PPML report a negative coefficient for common language and insignificant coefficient estimates for common colony and legal origin.

[22] Lagged effects of trade agreements are included in prior research. See, e.g., Baier and Bergstrand (2007), Bergstrand et al. (2015), and for pair-varying effects, see Baier et al. (2019). In our analysis, lag terms beyond five years and forward EIA effects are not significant.

[23] Equation (14) resembles the expressions for comparative statics in Behar and Nelson (2014). Two key differences exist, however. Our expressions do not distinguish extensive and intensive margins and

imply pair-specific comparative statics depending on the partners' weights s_i and s_j, and on the interaction terms z^k_{ijl}. Although we show them aggregated across all interactions, disaggregated measures can be calculated.

[24] Head and Mayer (2014) report a preferred estimate for the trade elasticity of 5.03, the median coefficient obtained using tariff variation, while controlling for multilateral resistance terms. Eaton and Kortum (2002) report an elasticity of 3.6 when instrumenting wages and 8.28 when using data on price gaps between countries to proxy for trade costs. Simonovska and Waugh (2011) refine the method of Eaton and Kortum (2002) based on price gaps and estimate an elasticity of 4.5. Arkolakis et al. (2012) report a preferred elasticity of 6.5 based on a Ricardian model.

[25] Modular, direct and indirect effects do not follow additivity, because they are calculated in exponential form.

[26] In Table 12 in Appendix D we compare results using different trade elasticities. Overall, increasing the tariff elasticity (in absolute terms) decreases the effects of EIAs and compresses the distribution of the equivalent trade friction changes. The histograms of the distribution of the equivalent trade friction changes for different trade elasticities can be obtained from the authors upon request.

[27] It should be noted that the measurement definitions do not match directly, as we report modular effects, while Baier et al. (2019) and Kohl (2014) report general-equilibrium effects, the difference being income and price effects. Although these effects may be large when they accumulate across pairs (see also Anderson et al. 2018), Baier et al. (2019), and Kohl (2014) use a relatively small sample of countries, such that the accumulation of income and price effects may not affect these shares dramatically.

[28] For a detailed picture of the evolution across all years in the sample, see the boxplots in Figure 4 in Appendix D.

[29] A simulation based on the model excluding the EU and NAFTA interactions (see column 4 in Table 1) produces significantly smaller estimated trade effects for the pairs in these agreements. However, other pairs are not affected and the effects across development groups are similar. Other differences between the analysis of Baier et al. (2018) and ours are the inclusion of internal trade, phase-in (lagged) EIA effects, the additional variables reflecting distribution costs in our analysis, and the method of estimation (OLS) in Baier et al. (2018). The difference in the method of estimation seems to underlie the difference in the results between our simulation and Baier et al. (2018). Using the parametrization based on PPML of Baier et al. (2018), the relationship between trade effects and GDPpc is positive.

[30] Up to 2004, 15 countries were part of the EU—Germany, France, Italy, The Netherlands, Belgium, Luxembourg, Denmark, Ireland, United Kingdom, Greece, Spain, Portugal, Austria, Finland, and Sweden. In 2004, Czech Republic, Estonia, Cyprus, Latvia, Lithuania, Hungary, Malta, Poland, Slovakia, and Slovenia joined the EU. Bulgaria and Romania did it in 2007, while Croatia joined in 2013.

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Supporting Information

Additional supporting information can be found online in the Supporting Information section.