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Economic integration agreements, border effects, and distance elasticities in the gravity equation

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ABSTRACT

Using a novel common econometric specification, we examine the measurement of three important effects in international trade that historically have been addressed largely separately: the (partial) effects on trade of economic integration agreements, international borders, and bilateral distance. First, recent studies focusing on precise and unbiased estimates of effects of economic integration agreements (EIAs) on members' trade may be biased upward owing to inadequate control for time-varying exogenous unobservable country-pair-specific changes in bilateral export costs (possibly decreasing the costs of international relative to intranational trade); we find evidence of this bias using a properly specified gravity equation. Second, our novel methodology yields statistically significant estimates of the declining effect of "international borders" on world trade, now accounting for endogenous EIA formations and unobserved country-pair heterogeneity in initial levels. Third, we confirm recent evidence providing a solution to the "distance-elasticity puzzle," but show that these estimates of the declining effect of distance on international trade are biased upward by not accounting for endogenous EIA formations and unobserved country-pair heterogeneity. We conclude our study with numerical general equilibrium comparative statics illustrating a substantive difference on trade effects of EIAs with and without allowance for the declining effects of international borders on world trade.

1. Introduction

It's a Small World After All…. (Walt Disney, New York World's Fair, 1964)

Using a novel common econometric specification, we examine the measurement of three important effects on international trade flows that have historically been addressed largely separately: the (partial) effects on trade of economic integration agreements (EIAs), international borders, and bilateral distance.[1] First, one of the most prominent aspects of the global economy over the past 20 years has been the proliferation of economic integration agreements (EIAs) – notably free

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[1] We are concerned primarily in this study with estimating partial (or direct) effects, not general equilibrium effects as in Anderson and van Wincoop (2003); Baier and Bergstrand (2009); Anderson and Yotov (2011), and Bergstrand et al. (2013). Nevertheless, general equilibrium effects hinge upon partial

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trade agreements but also some customs unions. Policy makers at national and supra-national government levels increasingly rely on *ex post* estimates of the (partial) effects of EIAs on trade flows based upon gravity equations to evaluate subsequently the welfare effects of EIAs.[2] Only recently have economists been able to provide more precise and unbiased *ex post* estimates of the effects of EIAs on members' international trade flows, in contrast to the highly variable and often economically implausible estimates generated over 45 years from 1962 to 2007.[3] Using panel data and accounting for the endogeneity of EIAs and prices and for unobserved country-pair heterogeneity, Baier and Bergstrand (2007), or BB, found using a sample spanning 1960–2000 and ordinary least squares (OLS) that a typical EIA increases two members' aggregate goods bilateral trade about 100 percent after 10–15years – five times the effect estimated using atheoretical gravity equations. Anderson and Yotov (2011) found similar results using the same BB specification (but using a Poisson quasi maximum likelihood (PQML) estimator) and showed the method also generated plausible, precise, and statistically significant effects for disaggregate trade flows. Eicher et al. (2012) use Bayesian Model Averaging (BMA) techniques to confirm the trade-creating effects of EIAs under model uncertainty. The key in BB, Anderson and Yotov (2011), and Eicher et al. (2012) was accounting for unobserved heterogeneity in exporters' and importers' time-varying multilateral influences (such as countries' prices and GDPs) and for unobserved heterogeneity in time-invariant bilateral influences. However, all of these studies failed to account for possible unobservable exogenous *time-varying* country-pair-specific changes in bilateral export costs (possibly decreasing the costs of international relative to intranational trade) that may have resulted in estimates of EIAs' effects being biased upward. The potential bias introduced by time-varying bilateral fixed export costs is especially important in light of their prominence in the "New New" trade theory, cf., Redding (2011) and Melitz and Redding (2014), and their empirical relevance, cf., Roberts and Tybout (1997) and Das et al. (2007). In this paper, we address this potentially important shortcoming using a properly specified gravity equation motivated by formal theoretical foundations. In doing so, we also contribute to two related literatures: the "(international) border puzzle" and the "distance-elasticity puzzle."

The "border puzzle" refers to the seminal estimate using traditional atheoretical gravity equations in McCallum (1995) that the Canada–U.S. international border caused Canadian inter-province trade to be 22 times – or 2100 percent – greater than province-state international trade in 1988, other things equal. This result implied that international borders imposed dramatic costs on international relative to intranational trade. This finding inspired an entire literature, including Anderson and van Wincoop's seminal (2003) paper formulating a new theoretical foundation for the gravity equation, building upon formal foundations in Anderson (1979) and Bergstrand (1985). While Anderson and van Wincoop (2003) addressed the importance of accounting properly for endogenous prices (in their terms, "multilateral resistances") in estimation and in general equilibrium comparative statics, to date estimates of the border effect are still very large. For instance, de Sousa et al. (2012) report that on average a country traded 493 times more intranationally than internationally in 1990 (cf., their Figure 1 for 1990 ($e^{6.2}=493$)), even dwarfing the McCallum estimate. Moreover, they estimate that on average this effect fell 63 percent to 181 in 2002 ($e^{5.2}=181$), that is, in only 12 years. However, using a time-series of cross-sections, they did not control for unobserved country-pair heterogeneity in border effects, did not account for endogenous EIAs as in BB, and, while recognizing multilateral prices in their estimation, did not account for the endogeneity of prices as addressed in Anderson and van Wincoop (2003). In this paper, we use an enhanced version of the BB panel-data methodology (and also an expansion of the data to include intranational trade flows) to provide consistent and precise estimates of the average declining effects of international borders on international trade, using a properly specified gravity equation accounting also for the effects of endogenous EIA formations, endogenous prices, and unobserved country-pair heterogeneity in initial border effects.

The "distance-elasticity puzzle" refers to the issue that – despite widespread anecdotal evidence that the effect of distance on international trade is declining over time, as suggested by Thomas Friedman's "flatter world" – systematic academic empirical evidence suggests that the distance elasticity of bilateral international trade has *not* declined, as established in the Disdier and Head (2008) meta analysis of the distance elasticity.[4] While some authors have offered alternative explanations, they have met mixed success (cf., footnote 1 in Yotov (2012) and our discussion later). However, Yotov (2012) recently provided a persuasive solution to the distance-elasticity puzzle by recognizing the importance of including *intranational*, alongside international, trade flows and bilateral distances in estimation (using OLS and PQML), a feature actually common to the "border effects" literature, cf., Wei (1996). By typically excluding intranational trade flows and intranational distances, gravity-equation estimates cannot identify the impact on international trade of international trade costs *relative to* intranational trade costs; previous studies of the distance-elasticity puzzle ignored this. However, Yotov (2012) suffered from two shortcomings. The study did not account for unobserved heterogeneity across country pairs and omitted controls for EIAs, potentially biasing upward his estimates of the declining effect of distance. Recently, Bosquet

(*footnote continued*)
effects, and we provide general equilibrium comparative statics of trade effects of EIAs with and without allowance for the declining effects of international borders on world trade at the end of the paper.

[2] See, for example, Berden et al. (2010) or, more recently, Head and Mayer (2014) and Costinot et al. (2014).

[3] In a meta-analysis of 1827 earlier studies (including several using flawed specifications), Cipollina and Salvatici (2010) find a range of estimates between 12 percent and 285 percent. Their mean effect is 80 percent and median effect is 46 percent.

[4] The international border puzzle differs from the distance-elasticity puzzle in the following respect. Typically, the border puzzle is associated with arguably economically implausible estimates of the *level* effect of an international border on international trade flows. By contrast, the distance-elasticity puzzle is only concerned with an absence of declines in the distance elasticity of international trade, not the average *level* of the distance elasticity *per se*.

and Boulhol (2015) using PQML included country-pair fixed effects to account for country-pair heterogeneity and exporter-time and importer-time fixed effects, but could not find declining distance elasticities; however, they did not include intranational trade and distances. In this paper, we address all these shortcomings and find statistically significant *declining* distance elasticities that indicate the upward bias in estimates in Yotov (2012).

This is the first paper to address all three related – but historically often disjointed – issues using a common econometric framework.[5] Using a state-of-the-art gravity equation with international and intranational trade flows, we reconcile methodologically and empirically all three issues. We provide three potential contributions. First, some bilateral fixed and variable export costs are unobservable. If some bilateral fixed and variable export costs are declining over time – as anecdotal evidence suggests – the Melitz (2003) model suggests that aggregate bilateral trade of existing exporters should expand (i.e., intensive margin response to lower variable trade costs), some previously domestic firms should select into exporting (i.e., extensive margin response to lower fixed and variable trade costs), and the number of domestic firms should decrease (due to market competition). Hence, omission of variables that account for some of the increase in international relative to intranational trade – other than an EIA – could bias upward EIA coefficient estimates under the approach in BB. Thus, the resounding "yes" claimed in BB in response to their question, "Do free trade agreements actually increase members' international trade?", may have been premature, as the issue of declining bilateral fixed and variable export costs was ignored.[6] In this paper, we use three alternative methods to account more fully for *time-varying* unobserved heterogeneity in bilateral international trade costs relative to intranational trade costs.[7] We find that previous estimates of EIAs' partial effects have been biased upward. When using the PQML estimator, the effects of an EIA fall by 30 percent.

Second, we draw upon the notion used in Anderson and van Wincoop (2003), or AvW, that international border dummies imbed international trade costs relative to intranational trade costs in gravity equations with international and intranational trade flows. In the cross-sectional context of AvW, their international border dummy variable measured 1 if two sub-national regions (Canadian province or U.S. state) were from different countries, and 0 otherwise (and hence from the same country). Thus, their international border dummy was an exogenous index of whether the trade flow was an international versus intranational flow. Here, we use a panel data set of international and intranational trade flows for a large number of country pairs for a large number of years. Since intranational trade is a nation's gross output less exports, we focus our analysis on manufactures trade, since exports are measured on a gross basis and data on manufactures gross output is available.[8] We then construct an exogenous dummy variable $INTERNATIONAL_{ij}$ that assumes the value 1 if the source (i) and destination (j) countries are different ($i \neq j$) and the value 0 if i and j are the same ($i=j$). By incorporating this variable interacted with a set of year dummies – creating $INTER_{ij,t}$, $INTER_{ij,t+1}$, etc. – and then using the BB panel approach, we can isolate the effect of EIAs on bilateral trade to determine how much an EIA actually increases two members' trade, but now accounting for any possibly declining trends in unobservable bilateral fixed and variable trade costs that may have increased international relative to intranational trade. Moreover, the coefficient estimates for the multiple $INTER_{ij,t}$ dummies also provide direct estimates of the changing (partial) effect of an international border on a pair's trade flow. A novel aspect of our approach – accounting for unobserved country-pair heterogeneity in a panel – is to allow the *level* of the international border effect to vary across every country pair using pair fixed effects (or pair fixed effects interacted with a trend); by contrast, previous studies constrain the level border effect to be identical across all country pairs or certain groups of country pairs using cross-sectional data (excluding country-pair fixed effects).[9] We find direct estimates of the (average) falling partial effect of an international border (after accounting for EIAs) using a specification motivated by a formal theoretical foundation for the gravity equation and avoiding the endogeneity bias (attributable to endogenous prices and endogenous EIA formations) present in several studies.[10] One of our estimates suggests that the cost of an international border (in terms of trade flows) has declined on average by 25.3 percent from 1990 to 2002, or about 2.4 percent per year.

Third, we will provide a battery of sensitivity analyses to determine the robustness of the results to phase-ins of agreements, lagged terms-of-trade effects, reverse causality, various estimation techniques, disaggregation, and accounting for firm-heterogeneity and country-selection biases introduced potentially by using aggregated data. In one sensitivity analysis, our alternative measure to our international border dummies to control for declining trade costs is an interaction of bilateral distance with year dummies. These results confirm the findings in Yotov (2012) that the effect of bilateral distance – owing to

[5] As we will clarify later, we note that studies of the border puzzle typically do not address the distance-elasticity puzzle, and vice versa. Also, some studies of the border puzzle and some of the distance-elasticity puzzle include dummy variables for economic integration agreements, but typically do not examine in detail how EIAs' effects are sensitive to the specifications. This paper addresses all three issues using a unified framework.

[6] Unfortunately, there is little systematic evidence of *observed* declines in bilateral fixed and variable export costs. A few studies have explored the issue of declining information technology (IT) costs, cf., Freund and Weinhold (2004), Tang (2006), and Berthelon and Freund (2008). However, all these studies include time-varying *multilateral*, not bilateral, indexes of IT factors. Such multilateral factors will be accounted for in our estimation using exporter-year and importer-year fixed effects. No study has yet accounted for time-varying declines in *bilateral* trade costs, except those associated with EIAs or currency unions.

[7] As will be discussed shortly, the three methods are interacting time-invariant pair fixed effects with a time trend, interacting a bilateral "international border" dummy with year effects, and interacting bilateral distance (and other standard gravity covariates) with year effects.

[8] We will address aggregate and disaggregate manufactures trade. In one robustness analysis, we will look at aggregate merchandise trade allowing GDP (a value added measure) to be an imperfect proxy for gross output. As we will discuss later, studies in the border-effects literature, starting with McCallum (1995) and Wei (1996), have incorporated intranational along with international trade flows.

[9] The level of a country-pair's border effect is imbedded in our fixed effect estimate.

[10] For instance, Head and Mayer (2000) and de Sousa et al. (2012) accounted for relative prices using measured national prices, but not for the endogeneity of prices as raised by Anderson and van Wincoop (2003).

likely falling bilateral variable and fixed export costs – *is* declining over time. Like Yotov (2012), we use a more appropriate measurement of intranational trade (using manufactures gross output and (gross) exports rather than aggregate trade and GDPs). However, unlike Yotov (2012) our specification accounts for unobserved time-varying bilateral heterogeneity and for the effects of EIA formations, which potentially biased upward Yotov's estimates of the declining effect of distance. We find evidence of this upward bias and estimate that the effect of distance on international trade has fallen by 1.2 percent per year.[11] Our results also suggest that the declining effect of international borders on trade and of distance on trade are two sides of the same coin; international trade costs have likely been declining relative to intranational trade costs.

The remainder of the paper is as follows. Section 2 provides theoretical background for the estimating equation that will be used. Section 3 addresses econometric issues and provides a data description. Section 4 discusses the empirical results for EIAs' partial effects on total bilateral manufactures trade flows, including the results of a series of sensitivity analyses, and provides estimates of the declining effect of international borders on trade. Section 5 discusses the results for disaggregate manufactures trade flows. Section 6 evaluates the sensitivity of the results to using aggregate *goods* trade flows, a broader sample of countries, and a longer time series. Section 7 uses an alternative method to account for declining bilateral export costs (other than EIAs), and provides estimates of the declining effect of distance on international relative to intranational trade. In Section 8, we distinguish quantitatively between the general equilibrium trade impacts of EIAs with *and* without allowance for declining effects of international borders on world trade. As for the case of partial effects, general equilibrium world trade impacts are significantly larger when the declining effects of distance are also accounted for. In Section 9, we conclude that the specifications suggested several years ago in Baier and Bergstrand (2007) to account for endogenous EIAs can be substantively improved by including international border dummies (or distances interacted with year dummies) in panel specifications to account for systematic declines over time in unobserved time-varying bilateral export costs (while *simultaneously* accounting for the heteroskedasticity bias in OLS estimates), including intranational trade and distance in samples, and accounting for unobserved time-invariant heterogeneity in country-pair international border level effects. In short, just as BB showed several years ago that panel techniques along with a properly specified gravity equation were critical to finding unbiased and precise EIA estimates, this paper is designed to show that an enhanced version of BB using panel techniques are critical to finding unbiased and precise estimates of EIA effects on international trade, of the declining effect of international borders on trade, and of the declining distance-elasticity of international trade.

2. Motivating the gravity-equation specification

The gravity equation has become the empirical workhorse for estimating partial effects of EIAs on members' trade flows. Arkolakis et al. (2012) demonstrated that a gravity equation surfaces for a large class of “quantitative trade models” that feature four main assumptions: (1) Dixit–Stiglitz preferences; (2) one factor of production (typically, labor); (3) linear cost functions; and (4) perfect or monopolistic competition.[12] Trade models satisfying these four assumptions are Armington (cf., Anderson and van Wincoop, 2003), Ricardian (cf., Eaton and Kortum, 2002), Krugman (1980), and Melitz (2003). Arkolakis et al. (2012) concluded that the gravity equation provides a common method for estimating the trade elasticity across these different approaches.

Arkolakis et al. (2012) show that a gravity equation consistent with all four theoretical frameworks in a K-country world can be described by

$$X_{ij,t} = \frac{\chi_{ij,t} N_{i,t} (w_{i,t}\tau_{ij,t})^{\epsilon} Y_{j,t}}{\sum_{k=1}^{K} \chi_{kj,t} N_{k,t} (w_{k,t}\tau_{kj,t})^{\epsilon}} \tag{1}$$

where $X_{ij,t}$ is the trade flow from exporter i to importer j in year t, $N_{i,t}$ is the measure of goods in country i that can be produced in year t, $w_{i,t}$ is the wage rate in country i, $\tau_{ij,t}$ is (gross) variable trade costs of country i's products into j, ϵ is the elasticity of imports with respect to variable trade costs ($\epsilon < 0$), $Y_{j,t}$ is aggregate expenditure in j, and $\chi_{ij,t}$ captures all “structural parameters” other than $\tau_{ij,t}$ (for example, fixed export costs from i to j, or $f_{ij,t}$).

For the purposes of this paper, the variables of interest are $\tau_{ij,t}, f_{ij,t}$, and $\chi_{ij,t}$. Typically, researchers have assumed that the formation of an EIA (such as a free trade agreement) between i and j lowers $\tau_{ij,t}$. However, EIAs are broad agreements reaching beyond elimination of tariff rates and variable trade costs; they likely also lower fixed export costs, $f_{ij,t}$. Yet, in reality, all types of “structural parameters” may alter $\chi_{ij,t}$ over time, such as informational considerations, historical factors, nontariff barriers unrelated to EIAs, and bilateral preferences, cf., Head and Mayer (2000). Thus, the use of *time-invariant* pair-specific fixed effects, as in BB, may be insufficient to isolate an unbiased partial effect of an EIA's formation on trade, via lowering $\tau_{ij,t}$, because trade flows may also be influenced by falling (rising) values of $f_{ij,t}$ ($\chi_{ij,t}$).

Moreover, every theoretical quantitative trade model yielding a gravity equation embodies “intranational trade” ($X_{ii,t}$), i.e., a country's domestic spending on its own products. Our novel approach in this paper – using international *and* intranational trade flows – is to introduce a variable $INTER_{ij,t}$ to account for average (across all pairs of different countries) declines in unobservable bilateral (fixed and variable) export costs (unassociated with EIAs), but in the context of a properly

[11] This contrasts with the stable estimates of the distance elasticity over time using PQML in Bosquet and Boulhol (2015), likely due to their omission of intranational trade and distances in their sample.

[12] Arkolakis et al. (2012) also note three macro-level restrictions: (1) trade is balanced; (2) aggregate profits are a constant share of aggregate revenues; and (3) the import demand system is constant elasticity of substitution (CES).

specified gravity equation motivated by a formal theoretical foundation. $INTER_{ij,t}$ is defined as the product of a year dummy, D_t, and a time-invariant binary variable, $INTERNATIONAL_{ij}$, which assumes the value 1 if the source and destination countries, i and j, respectively, are different countries ($i \neq j$) and the value 0 if i and j are the same country ($i = j$).

The economic motivation is explained most easily by reviewing first AvW.[13] Consider AvW's cross-sectional context of trade flows between and among Canadian provinces and U.S. states in 1993. In the context of AvW's Armington framework, trade costs were determined by two variables, bilateral distance ($DISTANCE_{ij}$) and a dummy for whether the two regions were in different countries ($=1$, if $i \neq j$) or the same country ($=0$, if $i = j$). In their paper, they used non-linear least squares to estimate their gravity equation to account for endogenous non-linear multilateral price terms. However, most researchers, such as Feenstra (2004), have focused since then on consistent estimation of the bilateral border dummy and bilateral distance coefficient estimates using a specification such as

$$\ln X_{ij} = \beta_0 + \beta_1 \ln DISTANCE_{ij} + \beta_2 INTERNATIONAL_{ij} + \eta_i + \theta_j + \upsilon_{ij}. \tag{2}$$

where η_i is an exporter fixed effect, θ_j is an importer fixed effect, and υ_{ij} is an error term. In AvW's cross-sectional context, $INTERNATIONAL_{ij}$ captures any factor influencing international relative to intranational trade. An international border imposes considerable costs. Thus, $INTERNATIONAL_{ij}$ would capture any cross-sectional variation in *bilateral* trade costs, beyond the role of bilateral distance which is also present in Eq. (2); $INTERNATIONAL_{ij}$ should have a negative effect on trade flows.[14] Of course, many other factors influence bilateral flows (international or intranational), so it is feasible to replace $INTERNATIONAL_{ij}$ and $\ln DISTANCE_{ij}$ with a country-pair fixed effect (in a panel). This is a novel aspect of our panel approach as previous border-effect studies have *not* used country-pair fixed effects to allow for variation across pairs in the *level* border effect.[15]

Our approach is to estimate each of the following equations using a panel of international and intranational trade flows:

$$X_{ij,t} = \exp\,[\beta_0 + \beta_1 EIA_{ij,t} + \beta_2 INTER_{ij,t} + \beta_3 DIST_{ij,t} + \eta_{i,t} + \theta_{j,t} + \phi_{ij}] + \zeta_{ij,t}, \tag{3}$$

and alternatively

$$X_{ij,t} = \exp\,[\beta_0 + \beta_1 EIA_{ij,t} + \beta_2 INTER_{ij,t} + \beta_3 DIST_{ij,t} + \eta_{i,t} + \theta_{j,t} + (\phi_{ij} \times Trend)] + \zeta_{ij,t}, \tag{4}$$

where $EIA_{ij,t}$ is a dummy assuming the value 1 (0) if an EIA exists (does not exist) in year t between countries i and j. Let *exp* denote the exponentiated value of the term in brackets and, for now, we allow the error term, $\zeta_{ij,t}$, to enter additively. This specification allows for estimation in levels using Poisson estimators and allows for zeros in trade.[16] In Eqs. (3) and (4), $\eta_{i,t}$ captures all time-varying multilateral factors of exporting country i, such as – in the context of Eq. (1) above – $N_{i,t}$ and $w_{i,t}$. $\theta_{j,t}$ captures all time-varying multilateral factors of importing country j, such as $Y_{j,t}$ and the denominator of the relative price term in Eq. (1).[17] Accordingly, our paper addresses the issue raised in AvW that direct inclusion of price variables creates potential endogeneity bias.[18] The inclusion of time-invariant country-pair fixed effects ϕ_{ij} captures all time-invariant factors that might otherwise be picked up by $EIA_{ij,t}$. However, there may exist trends over time in the effects of unobserved bilateral heterogeneity. Consequently, in Eq. (4) we interact the bilateral fixed effects (ϕ_{ij}) with a time trend (*Trend*).[19]

Yet, as Eq. (1) reveals, $\tau_{ij,t}$ and $\chi_{ij,t}$ are time-varying and reflect both policy-based and "natural" trade costs influencing international relative to intranational trade. However, there is a way to account for time-varying changes in these bilateral export costs, separate from policy-based trade liberalizations such as EIA formations. Recall that the variable $INTER_{ij,t}$ is defined as $INTER_{ij,t} = D_t \times INTERNATIONAL_{ij}$, where D_t is a year dummy. The variable $DIST_{ij,t}$ is defined analogously as the product of $\ln DISTANCE_{ij}$ and a year dummy.[20] In the presence of time-invariant pair fixed effects ϕ_{ij}, variation in $INTER_{ij,t}$ will capture all bilateral factors influencing international relative to intranational trade over time *on average* relative to the base period (hence, deviations over time relative to the pair fixed effect). Thus, any time-varying pair-specific variable such as $EIA_{ij,t}$ will capture only the effects on trade over time associated with the EIA's formation and not other factors causing $\tau_{ij,t}$ and $\chi_{ij,t}$ to change over time.

[13] An excellent treatment of this model and these issues is in chapter 5 of Feenstra (2004).

[14] Typically, gravity equations include other variables, such as dummies for sharing a common language, a common colonial history, or a common land border. Our empirical specifications below will account for these as well; we ignore them in Eqs. (2)–(4) to avoid clutter and abbreviate the equations in the text.

[15] It is important to note that some previous border-effect studies appropriately accounted for intranational as well international trade flows, such as Wei (1996), Head and Mayer (2000), Fontagne et al. (2005), and de Sousa et al. (2012). However, all such studies used a cross-sectional approach (for multiple years), including typical bilateral variables such as distance but constraining all country-pairs to have the same border-effect level. Our panel approach using country-pair fixed effects (and alternatively such effects interacted with a time trend) accounts for unobserved heterogeneity across country-pairs in initial border-effect levels. Moreover, by using our panel approach, we also account for the endogeneity of EIAs. Finally, all four studies noted above included variables representing prices, but did not account for the endogeneity of prices, i.e., the AvW critique. Wei (1996) included a linear approximation of the two countries' multilateral prices, but the approximation was not complete because it used only bilateral distances, cf., Baier and Bergstrand (2009).

[16] We address these econometric issues later.

[17] It is important to note that reductions in MFN tariff rates in importer j would be accounted for by $\theta_{j,t}$ as well.

[18] See footnote 15.

[19] Note that inclusion of an ij fixed effect *for each year* is infeasible; it would perfectly predict trade flows.

[20] As shown below, we will construct isomorphic dummies for common language, common colonial history, and common land border (i.e., contiguity). For brevity, we ignore these additional dummies in Eqs. (3) and (4).

Consequently, the addition of $INTER_{ij,t}$ alongside incorporating *intranational* trade flows and distances – consistent with theoretical foundations for the gravity equation explaining intranational as well as international flows – to the otherwise similar specifications in BB will essentially "purge" the partial EIA effects estimated in BB of omitted variables bias caused by changes in $\tau_{ij,t}$ and $\chi_{ij,t}$ unassociated with EIAs.[21] Moreover, the country-pair fixed effect ϕ_{ij} will capture the average trade-depressing effect of an international border, allowed to vary across all pairs. If trade costs unassociated with EIAs are falling over time, raising international relative to intranational trade, $INTER_{ij,t}$ will have a positive coefficient estimate. Also, the inclusion of $DIST_{ij,t}$ will account for any possible *additional* changing influence of distance over time on trade flows, not already captured by $INTER_{ij,t}$. As Yotov (2012) suggests, one of the factors most likely to explain increasing international relative to intranational trade is the declining effect of distance. However, that effect is likely to be absorbed by $INTER_{ij,t}$. In the absence of $INTER_{ij,t}$, $DIST_{ij,t}$ is likely to play a prominent role; we address this issue separately in a later section of the paper.

3. Econometric issues and data description

3.1. Econometric issues

The previous section dealt with many specification issues. However, one issue omitted above was the estimation approach. Historically, gravity equations have been estimated using OLS. The original analysis of the Canadian–U.S. "border puzzle" in McCallum (1995) used OLS. Baier and Bergstrand (2007) used OLS. Because of the introduction of a two-equation "structural" gravity model in AvW where one of the equations was non-linear, AvW used non-linear least squares. But with most of the literature focusing first on estimating the partial (or direct) effect of a border, most cross-section estimates have used OLS employing exporter and importer fixed effects (cf., Feenstra, 2004) and recently panel estimates have used OLS employing exporter-year, importer-year, and country-pair fixed effects, cf., Baier et al. (2014).

However, a large sub-literature of the gravity equation, starting with Haveman and Hummels (2004), Felbermayr and Kohler (2006), and Helpman et al. (2008), has addressed the importance of zeros in international trade flows, cf., Head and Mayer (2014). Santos Silva and Tenreyro (2006) showed using empirical specifications and a Monte Carlo analysis that, even in the absence of zeros, log-linear estimates of gravity equations suffered from heteroskedasticity bias (owing to Jensen's inequality). They showed that the PQML estimator could eliminate this heteroskedasticity bias as well as allow for inclusion of zeros. Consequently, as in many recent studies, we prefer the PQML estimator for Eqs. (3) and (4) above. However, we will show that our results are qualitatively the same using OLS.[22]

Summarizing, the key features of our specifications, for which previous analyses have excluded at least one dimension, are

1. Exporter-year and importer-year fixed effects to account for endogenous prices and unobserved time-varying exporter and importer multilateral heterogeneity;
2. Country-pair fixed effects or country-pair fixed effects interacted with a time trend to account for unobserved time-invariant or time-varying, respectively, bilateral effects, including pair-specific initial border effect levels;
3. Intranational as well as international trade flows and bilateral distances, so that the international border dummies can account for average declining international *relative to* intranational bilateral trade costs; and
4. PQML estimation to account for heteroskedasticity bias (owing to Jensen's inequality) and to account for zero trade flows.

3.2. Data description

Unlike the original estimates in BB, which examined aggregate bilateral goods trade flows, our analysis here focuses on manufactures trade flows. The reason is that a key right-hand-side (RHS) variable, $INTER_{ij,t}$, captures the effect *over time* of the (likely declining) average cost of international relative to intranational trade. Hence, as in McCallum (1995) and AvW, the left-hand-side (LHS) variable needs to include observations on *intranational* trade. Since exports are measured on a "gross" (not value added) basis, national output needs to be measured on a comparable basis to estimate intranational trade. The data used are the sectoral manufacturing data from Anderson and Yotov (2011). These data cover 41 trading partners (40 separate countries and a Rest-of-World (*ROW*) aggregate, consisting of 24 additional nations).[23] The eight manufacturing

[21] BB actually used country-and-year dummies (rather than exporter-and-year and importer-and-year dummies) along with pair fixed effects. The specifications here are akin to Baier et al. (2014), which used exporter-and-year, importer-and-year, and pair fixed effects.

[22] See the recent survey of the gravity equation literature in Head and Mayer (2014) for a useful discussion of new directions on these estimation issues. See also SantosSilva and Tenreyro (2011) on the robustness of PQML.

[23] Results will be provided with and without the *ROW* aggregate. The 40 main countries are Argentina, Australia, Austria, Belgium–Luxembourg, Bolivia, Brazil, Bulgaria, Canada, Chile, China, Colombia, Costa Rica, Denmark, Ecuador, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Korea (South), Mexico, Morocco, The Netherlands, Norway, Poland, Portugal, Romania, Spain, Sweden, Switzerland, Tunisia, Turkey, United Kingdom, United States, and Uruguay. The 24 countries in the *ROW* aggregate are Cameroon, Cyprus, Egypt, Hong Kong, Indonesia, India, Iran, Jordan, Kenya, Kuwait, Sri Lanka, Macao, Malta, Myanmar, Malawi, Malaysia, Niger, Nepal, Philippines, Senegal, Singapore, Trinidad and Tobago, Tanzania, and South Africa.

sectors are classified according to the United Nations' 2-digit International Standard Industrial Classification (ISIC) Revision 2.[24] The period of investigation is 1990–2002. For our analysis, we use only the years 1990, 1994, 1998, and 2002, akin to BB's use of data for every five years. We use a shorter four-year interval than BB's five-year interval due to the shorter time-series for our data. However, the use of every four years (or five years in BB) addresses the concern raised in Cheng and Wall (2005) that "Fixed-effects estimations are sometimes criticized when applied to data pooled over consecutive years on the grounds that dependent and independent variables cannot fully adjust in a single year's time" (p. 8). Also, Wooldridge, 2009 confirms the reduction in standard errors of coefficient estimates using changes over longer periods of time than using "year-to-year" changes (p. 459).[25]

Bilateral international trade flows are defined as the value of exports from exporter i to importer j. We use the CEPII *Trade, Production and Bilateral Protection Database*[26] (TradeProd) as the main trade data source because it implements a consistent procedure for mapping the CIF (cost, insurance and freight) values reported by the importing countries in COMTRADE to the FOB (free on board) values reported by the exporters in COMTRADE.[27] This decreases the number of missing observations in the sample.[28] To further decrease the number of missing trade flows, we add export values from the United Nations Statistical Division (UNSD) Commodity Trade Statistics Database (COMTRADE).[29] Internal commodity-level (intranational) trade for each country is constructed as the difference between total manufactures output and aggregate manufactures exports to all trading partners, which come from the same data sources. The number of zero trade flows in the sample is very small and we will document this later. This suggests that the consequences of throwing information away by using the standard log-linear OLS estimator should not be severe. Nonetheless, the PQML estimator is still preferable because, in addition to accounting for the zero trade flows, it also controls for heteroskedasticity bias introduced due to Jensen's inequality.

Industrial gross output value data come from two sources. The primary source is the United Nations' UNIDO Industrial Statistics database, which reports industry-level output data at the 3-digit and 4-digit level of ISIC Code (Revisions 2 and 3). We use the CEPII TradeProd database as a secondary source of product-level output data.[30] We interpolate some of the missing output values for the sample countries, which account for 15.6% of the observations; however, results are robust to inclusion or exclusion of observations using interpolated output values.

Data on EIA dummies come from the Database on Economic Integration Agreements on Jeffrey Bergstrand's website (http://www.nd.edu/ jbergstr). Baier and Bergstrand's EIA database categorizes bilateral EIA relationships from 1950–2005 for pairings of 195 countries using a multichotomous index. In this study, $EIA_{ij,t}=1$ denotes a free trade agreement between a pair of countries ij in year t or deeper integration (such as a customs union, common market, or economic union), or 0 otherwise, as in BB. Table 1 lists the agreements. In our *ROW* aggregate, there are no countries with EIAs with the main 40 countries. Initially, our results will include observations for $ROW_{ij,t}$; however, in a sensitivity analysis, we show the results are insensitive to the inclusion or exclusion of observations for $ROW_{ij,t}$ or for those with interpolated output values.[31]

Following Mayer and Zignago (2006), bilateral distance (between countries and internally within countries) is calculated as $DISTANCE_{ij}=(\sum_{k\in i}Pop_k/Pop_i)\ (\sum_{l\in j}Pop_l/Pop_j)D_{kl}$, where Pop_k is the population of agglomeration k in country i in year 2004, Pop_l is the population of agglomeration l in country j, and D_{kl} is the bilateral distance in kilometers between agglomeration k and agglomeration l (using Great Circle Distance formula). All data on latitudes, longitudes, and population are from the World Gazetteer web page. An important feature of this variable is that the same procedure is used to construct (consistently) international as well as intranational distances. The other dummy variables in our analysis are from CEPII.

4. Empirical results for total manufacturing trade flows

Table 2 presents our main results using aggregate international and intranational manufactures bilateral trade flows and some alternative specifications for robustness. Table 2 is partitioned into two panels, 2A and 2B. Panel 2A provides estimates omitting our key variable $INTER_{ij,t}$, as well as $DIST_{ij,t}$ and the analogous dummies for sharing a common language ($LANG_{ij,t}$), a

[24] The nine 2-digit ISIC manufacturing categories are (short labels, used for convenience throughout the paper, are noted in parentheses): 31. Food, Beverages, and Tobacco Products (Food); 32. Textile, Apparel, and Leather Products (Textile); 33. Wood and Wood Products (Wood); 34. Paper and Paper Products (Paper); 35. Chemicals, Petroleum, Coal, Rubber, and Plastic Products (Chemicals); 36. Other Non-metallic Products (Minerals); 37. Basic Metal Industries (Metals); 38. Fabricated Metal Products, Machinery, Equipment (Machinery); 39. Other manufacturing. Inspection of the output data at the 3-digit and 4-digit ISIC level of aggregation reveals that many countries report Equipment production, and especially Scientific Equipment production, under the category Other Manufacturing. Therefore, to avoid inconsistencies, we combine the last two 2-digit categories into one, which we label Machinery.

[25] For robustness, we will also examine later the sensitivity of our findings to aggregate goods trade flows and GDPs, using a broader sample of countries representing 98 percent of world output and for a longer time period.

[26] For details regarding this database see Mayer et al. (2008).

[27] The TradeProd database is based on the CEPII *Base pour l'Analyse du Commerce International* (BACI) data. For details regarding BACI see Gaulier and Zignago (2008).

[28] As noted in Anderson and Yotov (2010), in principle, gravity theory calls for valuation of exports at delivered prices. In practice, valuation of exports FOB avoids measurement error arising from poor quality transport cost data.

[29] We access COMTRADE through the World Integrated Trade Solution (WITS) software, http://wits.worldbank.org/witsweb/ The software reports trade data in three different concordances including Harmonized System (HS) Revisions 1989/92 and 1996, and the Standard International Trade Classification (SITC), which are automatically converted to ISIC Rev. 2. To obtain maximum number of observations, we combine the data from the different concordances.

[30] TradeProd uses the OECD STAN Industrial Database in addition to UNIDO's Industrial Statistics Database.

[31] An online appendix reports results excluding observations using the *ROW* aggregate region and observations using interpolated output values.

Table 1
Economic Integration Agreements in the Data Set.

European Union, or EU (1958): Belgium–Luxembourg, France, Italy, Germany, The Netherlands, Denmark (1973), Ireland (1973), United Kingdom (1973), Greece (1981), Portugal (1986), Spain (1986), Iceland (1994) Austria (1995), Finland (1995), Sweden (1995)
European Free Trade Association, or EFTA (1960): Austria (until 1995), Denmark (until 1973), Iceland (1970), Finland (1986–1995), Norway, Portugal (until 1986), Sweden (until 1995), Switzerland, United Kingdom (until 1973)
Latin American Free Trade Agreement/Latin American Integration Agreement, or LAFTA/LAIA (1993–): Argentina, Bolivia, Brazil, Chile, Ecuador, Mexico, Uruguay
EU–EFTA Agreement/European Economic Area (1973/1994)
US–Israel (1985)
US–Canada (1989)
EFTA–Israel (1993)
Central Europe Free Trade Agreement, or CEFTA (1993): Hungary, Poland, Romania (1997), Bulgaria (1998)
EFTA–Turkey (1992)
EFTA–Bulgaria (1993)
EFTA–Hungary (1993)
EFTA–Poland (1993)
EFTA–Romania (1993)
Andean Community (1993): Bolivia, Columbia, Ecuador
EU–Hungary (1994)
EU–Poland (1994)
North American Free Trade Agreement, or NAFTA (1994): Canada, Mexico, United States
Bolivia–Mexico (1995)
Costa Rica–Mexico (1995)
EU–Bulgaria (1995)
EU–Romania (1995)
Columbia–Mexico (1995). As part of the Group of Three. The third country, Venezuela, is not in the sample.
Mercosur (1991): Argentina, Brazil, Uruguay (formed in 1991 FTA in 1995)
Mercosur–Chile (1996)
Mercosur–Bolivia (1996)
EU–Turkey (1996)
Canada–Chile (1997)
Canada–Israel (1997)
Hungary–Turkey (1998)
Hungary–Israel (1998)
Israel–Turkey (1998)
Romania–Turkey (1998)
Poland–Israel (1998)
EU–Tunisia (1998)
Mexico–Chile (1999)
EU–Israel Agreement (2000)
EU–Mexico (2000)
EU–Morocco (2000)
EFTA–Morocco (2000)
Poland–Turkey (2000)
Mexico–Israel (2000)
Chile–Costa Rica (2002)

Notes: This table lists, in chronological order, all economic integration agreements (EIAs) used in estimation. Only agreements involving the countries in our sample are included. EIAs that entered into force before 1990 are used, as needed, to construct the lagged variables of the EIA dummy variable; the EIA data set includes EIA dummies from 1950 to 2011. For all estimates, EIAs include free trade agreements and deeper integration agreements based upon the Baier–Bergstrand data set.

common colonial history ($CLNY_{ij,t}$), and a common land border or contiguity ($CNTG_{ij,t}$). Panel 2B includes $INTER_{ij,1994}$, $INTER_{ij,1998}$, and $INTER_{ij,2002}$; $INTER_{ij,1990}$ is omitted due to the inclusion of a constant.[32]

The first specification for column (1) in panel 2A is based upon Eq. (3) using PQML. We include lagged values of $EIA_{ij,t}$ along with the current value, based upon numerous previous studies that have found effects lasting up to 10–15 years. Typically, most EIAs are phased in over 5–10 years. Moreover, EIAs likely have lagged effects on trade as flows respond slowly to terms-of-trade changes. Lagged values of EIAs capture these influences. As discussed above, our preferred econometric estimator is PQML; however, later in Table 2 we present results using OLS. To show that the zeros issue is not a problem, our first specification is on positive trade flows only. Consistent with previous findings, column (1) in panel 2A indicates that the partial effects of EIAs are economically and statistically significant, both for current and most lagged values of $EIA_{ij,t}$. The total (partial) effect of an EIA is 157 percent ($=100\times[e^{0.945}-1]$). Compared to earlier studies using OLS and aggregate goods trade flows this estimate is large; however, our study uses PQML and manufactures trade flows. As first

[32] However, due to the inclusion of the other fixed effects, the constant cannot be interpreted as an estimate of $INTER_{ij,1990}$.

Table 2
Estimates for manufactures trade flows.

2A. EIAs	(1) PQML(+)	(2) PQML	(3) PQML Lead	(4) PQML Trend	(5) PQML sub-sample	(6) OLS
$EIA_{ij,t}$	0.241	0.241	0.239	0.251	0.237	0.132
	(0.038)**	(0.038)**	(0.048)**	(0.039)**	(0.037)**	(0.055)*
$EIA_{ij,t-4}$	0.294	0.294	0.294	0.299	0.292	0.123
	(0.050)**	(0.050)**	(0.049)**	(0.050)**	(0.053)**	(0.051)*
$EIA_{ij,t-8}$	0.248	0.248	0.248	0.253	0.258	0.025
	(0.033)**	(0.033)**	(0.034)**	(0.034)**	(0.032)**	(0.051)
$EIA_{ij,t-12}$	0.069	0.069	0.069	0.074	0.066	0.027
	(0.045)	(0.045)	(0.045)	(0.046)	(0.045)	(0.063)
$EIA_{ij,t-16}$	0.093	0.093	0.093	0.100	0.094	0.014
	(0.030)**	(0.030)**	(0.030)**	(0.030)**	(0.031)**	(0.068)
$EIA_{ij,t+4}$			0.005			
			(0.048)			
Total EIA	0.945	0.944	0.944	0.977	0.948	0.321
	(0.085)**	(0.085)**	(0.095)**	(0.088)**	(0.084)**	(0.141)*
N	6639	6724	6724	6724	6116	6639
2B. EIAs and INTER						
$EIA_{ij,t}$	0.111	0.110	0.144	0.111	0.114	0.110
	(0.043)*	(0.043)*	(0.047)**	(0.043)**	(0.045)*	(0.051)*
$EIA_{ij,t-4}$	0.218	0.218	0.215	0.216	0.222	0.114
	(0.041)**	(0.041)**	(0.039)**	(0.041)**	(0.044)**	(0.050)*
$EIA_{ij,t-8}$	0.159	0.159	0.155	0.157	0.178	0.017
	(0.029)**	(0.029)**	(0.029)**	(0.029)**	(0.032)**	(0.050)
$EIA_{ij,t-12}$	−0.024	−0.024	−0.023	−0.026	−0.023	0.010
	(0.027)	(0.027)	(0.027)	(0.028)	(0.026)	(0.066)
$EIA_{ij,t-16}$	0.044	0.044	0.043	0.045	0.057	0.002
	(0.030)	(0.030)	(0.030)	(0.031)	(0.031)+	(0.068)
$EIA_{ij,t+4}$			−0.089			
			(0.052)+			
$INTER_{ij,1994}$	0.099	0.099	0.104	0.107	0.061	0.387
	(0.021)**	(0.021)**	(0.022)**	(0.021)**	(0.021)**	(0.028)**
$INTER_{ij,1998}$	0.282	0.282	0.286	0.298	0.259	0.654
	(0.028)**	(0.028)**	(0.030)**	(0.029)**	(0.033)**	(0.036)**
$INTER_{ij,2002}$	0.292	0.292	0.296	0.316	0.250	0.689
	(0.043)**	(0.043)**	(0.044)**	(0.044)**	(0.055)**	(0.044)**
Total EIA	0.507	0.506	0.534	0.504	0.550	0.252
	(0.099)**	(0.099)**	(0.102)**	(0.100)**	(0.114)**	(0.135)+
N	6639	6724	6724	6724	6116	6639

Notes: This table reports panel gravity estimates with data on total manufacturing, 1990–2002. All specifications include exporter-time, importer-time and country-pair fixed effects. Fixed effects estimates are not reported for brevity. Robust standard errors, clustered by country pair, are in parentheses. Panel 2A offers different variations of the main specification from Baier and Bergstrand (2007). In Panel 2B, we account for declining international border effects using time-varying border variables $INTER_{ij,t}$. Column 1, PQML(+), presents PQML estimates using only positive trade flows. The estimates in column (2), PQML, use all observations in the sample. In column (3), PQML Lead, we test for reverse causality by introducing a lead EIA effect. The estimates in column (4), PQML Trend, are obtained with pair-fixed effects interacted with a time trend. In column (5), PQML sub-sample, we drop the observations for the Rest-of-World (*ROW*) region and the observations that rely on interpolated output data. Finally, column (6) reports OLS estimates. See text for further details.
+ $p < 0.10$
* $p < .05$
** $p < .01$

seen in Santos Silva and Tenreyro (2006), PQML estimates tend to be systematically larger than OLS estimates for EIAs' coefficients. Also, highly tradable manufactures tend to be impacted much more by EIAs than most non-manufactured goods; evidence will be provided later by contrasting results for manufactures trade with those for aggregate goods trade.

The first specification in column (1) in panel 2B, also based on Eq. (3), includes $INTER_{ij,t}$. The first notable finding is a substantive decline in column (1) from panel 2A to panel 2B in the total (partial) EIA effect from 0.945 (157 percent) to 0.507 (66 percent). The impact of a typical EIA is *halved*. The second notable finding is that the effect of an international border has been declining over time, and this effect is economically and statistically significant. The coefficient estimate of 0.292 for $INTER_{ij,2002}$ implies that the declining effect of the international border has increased international relative to intranational trade by 34 percent over 12 years ($=100\times[e^{0.292}-1]$), or approximately 2.5 percent per year.[33] Hence, the costs of

[33] de Sousa et al. (2012) only report declining border effects using OLS. Bosquet and Boulhol (2015) estimate the changing effects of distance and common land borders using PQML; however, as discussed later, they do not find falling distance elasticities, likely due to omitting intranational trade. Note

international borders on international relative to intranational trade have been declining over time about 2.4 percent per year ($=100 \times [1-(1/e^{0.292})^{1/12}]$) and the absence of accounting for international border effects biases upwards EIA effects.

The second specification in panels 2A and 2B is identical to that in column (1) except that zeros are added to the sample. The addition of zeros to the sample has no material bearing on our results, partly due to the small number of zeros. However, for completeness we include them in our remaining specifications.

The specification in column (3) considers the potential bias created by "reverse causality." Wooldridge (2010, p. 325) suggests that it is easy to test for the "strict exogeneity" of EIAs in our context. To do this, we add a future level of EIA to the model. In the panel context here, if EIA is exogenous to trade flows, $EIA_{ij,t+4}$ should be uncorrelated with the current trade flow. Column (3) in panels 2A and 2B reports the results. In panel 2A, the future level of EIA has no economically or statistically significant effect on current trade. In Panel 2B, the effect is only marginally economically and statistically significant. Consequently, we do not consider reverse causality to be a serious problem.

The specification in column (4) is based upon Eq. (4). This specification allows for a trend change in the effect of the pair fixed effects, cf., (Wooldridge, 2010). Note that inclusion of *ij* fixed effects interacted with year dummies is infeasible; it would perfectly predict trade flows. Column (4) reports the results using Eq. (4), which are directly comparable to those in column (2). It is evident that there is not much difference in the results using equation specifications (3) or (4). Consequently, for the remainder of the analysis we focus on Eq. (3) specifications, which are computationally less demanding.

The specification in column (5) is based upon Eq. (3), but excludes observations for *ROW* and for any observations using interpolated output values. The specification in (5) is identical to that in column (2). Importantly, in both panels the EIA coefficient estimates are essentially the same; there are neither economically nor statistically significant differences. The only noticeable difference in panel 2B between columns (2) and (5) is that the coefficient estimates for the $INTER_{ij,t}$ dummies are larger in column (2) relative to (5). However, even by 2002, the declining effect of the international border on raising international relative to intranational trade is only 6 percent more in column (2) relative to column (5) ($[e^{0.292}-e^{0.250}] \times 100$). Moreover, none of these differences is statistically significant. Hence, for the remainder of the analysis we include observations for *ROW* and the observations using interpolated output values; however, the results excluding those are available in an online appendix.

While PQML has desired econometric properties as discussed above, the literature has not accepted completely a preference for PQML, cf., Head and Mayer (2000). Much earlier work in this literature has used OLS. For completeness, we also run the specification based upon Eq. (3) using OLS for positive trade flows only. Two results, reported in column (6) and comparing to columns (1) or (2), are worth noting. First, as found in most other studies, EIA effects are smaller using OLS compared to those using PQML, cf., Santos Silva and Tenreyro (2006). The total EIA effect in panel 2B using OLS is 0.252 relative to the PQML estimate of 0.507. Second, using OLS the coefficient estimates for $INTER_{ij,t}$ are higher.[34]

We provide one more sensitivity analysis using total manufactures trade flows. Reflecting upon Eq. (3), this specification suggests that – in principle – there is no reason not to include also variables analogous to $INTER_{ij,1994}$, $INTER_{ij,1998}$, and $INTER_{ij,2002}$, such as for bilateral distance and dummies for common language, common colonial history, and contiguity. Practically, a potential concern is that – since variables such as bilateral distance are highly correlated with a dummy for international relative to intranational trade – their additional presence may cause a multicollinearity problem. Nevertheless, we expanded the main specification in column (2) in panel 2B to also include $DIST_{ij,1994}$, $DIST_{ij,1998}$, $DIST_{ij,2002}$ (defined above) and analogous variables for common language ($LANG_{ij,t}$), common colonial history ($CLNY_{ij,t}$), and contiguity ($CNTG_{ij,t}$). These results are reported in column (3) of Table 3. For comparison, Table 3 first provides in column (1) a specification identical to that in Table 2, column (2), panel A. The total EIA (partial) effect is 0.944. In column (2) of Table 3, we add the $INTER_{ij,t}$ variables; as in the identical specification in Table 2, column (2), panel B, the total EIA effect falls significantly as expected. In column (3) of Table 3, we add $DIST_{ij,1994}$, $DIST_{ij,1998}$, $DIST_{ij,2002}$ and the analogous variables for common language ($LANG_{ij,t}$), common colonial history ($CLNY_{ij,t}$), and contiguity ($CNTG_{ij,t}$). Two main conclusions result. First, these additional variables – with two exceptions – have no significant effects on trade flows; the two exceptions are $CNTG_{ij,1994}$ and $CNTG_{ij,2002}$. Second, the EIA effects are higher in column (3) relative to column (2); there is some muting of the decline in EIA effects due to the additional variables.

The explanation for the limited explanatory power of the standard gravity covariates is straightforward. The bilateral variables that most influence $INTER_{ij,t}$ – a dummy for international relative to intranational bilateral trade – are likely $DIST_{ij,t}$, $LANG_{ij,t}$, $CLNY_{ij,t}$, and $CNTG_{ij,t}$. The presence of $INTER_{ij,t}$ subsumes the effects of those other variables; this leads to low coefficient estimates for the additional variables. Nevertheless, in a later sensitivity analysis that will help resolve the "distance-elasticity puzzle," we will demonstrate that – in the *absence* of the $INTER_{ij,t}$ dummies – these additional variables (notably, $DIST_{ij,t}$) have significant roles. Finally, Table 3, column (4) reports the results for the Eq. (4) version of column (3)'s specification; there is no material difference between the results in columns (3) and (4). Since Eq. (3) is computationally less demanding, we will use Table 3's column (3) specification for the baseline rather than column (4)'s specification.

(*footnote continued*)
that the initial international border effect levels for each country pair in our study are allowed to differ across pairs, and are subsumed in the pair fixed effects.

[34] The OLS total partial EIA estimate of 0.321 in panel 2A is within the range of partial EIA effect estimates reported earlier, where the median is about 0.46. The lower estimate here reflects estimation over a more recent time period (1990–2002) than used in most studies, such as BB which used 1960–2000. As found in recent research related to Baier et al. (2015), EIA partial effects have declined over the decades since 1960 because the (likely) most welfare-improving EIAs formed earlier and had larger partial effects.

Table 3
Estimates for manufactures trade flows.

	(1) PQML	(2) PQML	(3) PQML	(4) PQML Trend
$EIA_{ij,t}$	0.241	0.110	0.117	0.117
	(0.038)**	(0.043)*	(0.044)**	(0.044)**
$EIA_{ij,t-4}$	0.294	0.218	0.243	0.244
	(0.050)**	(0.041)**	(0.043)**	(0.043)**
$EIA_{ij,t-8}$	0.248	0.159	0.225	0.226
	(0.033)**	(0.029)**	(0.035)**	(0.036)**
$EIA_{ij,t-12}$	0.069	−0.024	0.020	0.021
	(0.045)	(0.027)	(0.037)	(0.036)
$EIA_{ij,t-16}$	0.093	0.044	0.022	0.023
	(0.030)**	(0.030)	(0.035)	(0.035)
$INTER_{ij,1994}$		0.099	0.171	0.175
		(0.021)**	(0.074)*	(0.074)*
$INTER_{ij,1998}$		0.282	0.231	0.241
		(0.028)**	(0.093)*	(0.094)**
$INTER_{ij,2002}$		0.292	0.294	0.309
		(0.043)**	(0.145)*	(0.146)*
$DIST_{ij,1994}$			−0.024	−0.022
			(0.029)	(0.029)
$DIST_{ij,1998}$			0.035	0.039
			(0.037)	(0.038)
$DIST_{ij,2002}$			0.021	0.026
			(0.055)	(0.056)
$LANG_{ij,1994}$			0.032	0.032
			(0.048)	(0.048)
$LANG_{ij,1998}$			0.005	0.004
			(0.073)	(0.073)
$LANG_{ij,2002}$			0.026	0.025
			(0.075)	(0.075)
$CNTG_{ij,1994}$			−0.144	−0.145
			(0.065)*	(0.066)*
$CNTG_{ij,1998}$			−0.078	−0.080
			(0.094)	(0.095)
$CNTG_{ij,2002}$			−0.226	−0.229
			(0.112)*	(0.112)*
$CLNY_{ij,1994}$			−0.034	−0.034
			(0.045)	(0.045)
$CLNY_{ij,1998}$			−0.030	−0.030
			(0.071)	(0.071)
$CLNY_{ij,2002}$			−0.057	−0.057
			(0.070)	(0.070)
Total EIA	0.944	0.506	0.626	0.630
	(0.085)**	(0.099)**	(0.111)**	(0.111)**
N	6724	6724	6724	6724

Notes: This table reports panel gravity estimates with data on total manufacturing, 1990–2002. All specifications include exporter-time, importer-time and country-pair fixed effects. Fixed effects estimates are not reported for brevity. Robust standard errors, clustered by country pair, are in parentheses. Columns (1) and (2) reproduce the results from column (2) of panel 2A and from column (2) of panel 2B, respectively. Column (3) introduces the standard gravity covariates as described in the text. Finally, the results in column (4) are obtained with pair-fixed effects scaled up by a time trend. See text for further details.

+ $p < 0.10$.

* $p < .05$

** $p < .01$

One potential bias we have not accounted for is firm-heterogeneity bias. As discussed in Helpman et al. (2008), or HMR, and Egger et al. (2011), the existence of firm heterogeneity may bias coefficient estimates in gravity equations using aggregate data. One of the advantages of HMR's two-stage approach is that it accounts for zeros, but also for firm heterogeneity, when using aggregate trade flows. HMR concluded that firm-heterogeneity bias mattered even more than country-selection bias in their cross-section estimates.[35] However, accounting also for endogeneity (self-selection) bias of EIAs, Egger et al. (2011) found that firm-heterogeneity bias hardly mattered at all. We argue here that – for our panel

[35] HMR also estimated their model pooling several cross-sections over time. However, their estimation never included bilateral country-pair fixed effects in their second stage, which is critical to our discussion below.

specification shown in Eq. (3) – the results are not likely influenced materially by firm-heterogeneity bias, due to the inclusion of the bilateral country-pair fixed effects. This is an issue explored only recently in Baier et al. (2014).

To understand why, we first review briefly the HMR approach, which was used in a cross section (1986 trade flows). The two-stage methodology entails estimating first a probit equation to determine the probability of a positive observation between a country pair. The probit estimates are then used to construct inverse Mills' ratios (denoted $IMR_{ij,t}$) to capture selection bias and variables $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$ to control for heterogeneous productivities of firms. $IMR_{ij,t}$, $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$ are then used as additional regressors in the second-stage gravity-equation specification to control for country-selection and firm-heterogeneity biases.[36] Both stages of estimation in HMR used exporter and importer fixed effects to account for multilateral variables, but did not use bilateral country-pair fixed effects (as HMR used a cross-section).

In our panel specification of Eq. (3), we have time-invariant bilateral country-pair fixed effects. If most of the variation in the predicted probit probabilities of trading (i.e., selection of country-pairs into positive trade) is cross-sectional in nature, then time-invariant country-pair bilateral fixed effects in the second stage will account for most of the variation in $IMR_{ij,t}$, $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$. The only possible bias in gravity equation coefficient estimates using our panel attributable to selection and firm-heterogeneity would be time variation in $IMR_{ij,t}$, $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$. It becomes an empirical issue then to determine if such bias is material.

Baier et al. (2014), or BBF, recently addressed the HMR two-stage estimation procedure in a panel with bilateral fixed effects (and alternatively first-differencing) in the second stage. Akin to HMR, BBF first estimated eight individual cross-section probits for the years 1965, 1970,…, 2000 to generate predicted probabilities of positive aggregate goods trade flows for a large number of country pairs. They then used these predicted probabilities to construct for each year $IMR_{ij,t}$, $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$. In the second stage, they estimated a specification similar to Eq. (3), but excluding $INTER_{ij,t}$ and using OLS. Their results from the second stage regressions were reported in BBF's Appendix Table A4, which can be readily compared to the results from omitting $IMR_{ij,t}$, $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$ which were presented in Table 1 of BBF. A comparison of the results from the two tables reveals clearly that there is very little quantitative and *no* qualitative differences between the respective coefficient estimates. The reason is the presence in the second stage of the bilateral fixed effects (or first-differencing). Put simply, most of the variation in the predicted probabilities of positive trade flows is cross-sectional, not time-varying; bilateral fixed effects (or first-differencing) accounts largely for the influences of country-selection and firm-heterogeneity. Based upon those results, we argue our results are likely robust to firm-heterogeneity bias. Moreover, one feature of our data is that there are very few zero trade flows that are not perfectly explained by our fixed effects (country-pair, exporter-time, and importer-time). This implies that there are few country-pairs that start or stop trading in our data set.[37]

5. Empirical results for disaggregate manufactures trade

One dimension which BB ignored entirely is the sensitivity of the findings to disaggregation of trade flows. The empirical literature on partial effects of EIAs using disaggregate data is actually quite small. Anderson and Yotov (2011) is one of the few studies using the BB approach to analyze disaggregate trade flow effects, and our study allows us to explore disaggregation since it is based upon the same data. Table 4 provides the results of estimating the partial EIA effects using Eq. (3) for the eight 2-digit ISIC categories of manufactures. Table 4 is divided into two panels, 4A and 4B. Panels 4A and 4B provide the main PQML specification results using each of eight manufactures industries. In Table 4, *DIST*, *LANG*, *CLNY*, and *CNTG* are included for the reasons discussed above.

Panel 4A provides the results of EIA effects in the analysis of disaggregate trade flows using PQML and excluding $INTER_{ij,t}$, $DIST_{ij,t}$ and the analogous variables. The results are largely consistent with those in the previous section for total manufactures trade flows. First, compare the results in Panel 4A of Table 4 with those in column (1) of Table 3. Using the total EIA estimates, we find positive and economically and statistically significant partial effects of EIAs on trade flows in seven out of the eight sectors, with the total partial effects of these seven sectors ranging from 0.175 (Wood) to 1.641 (Textile). Only in the Wood sector we do not find significant total effects of EIAs.

Panel 4B provides the results including the international border dummies and associated gravity variables. As in previous estimates, the EIA effects tend to diminish, but many remain economically and statistically significant. Using the total EIA estimates again, the partial effect for Machinery Products falls from 1.145 to 0.640. Yet, this still implies that an EIA increases trade by 90 percent (absent any general equilibrium effects). Estimates of the sums of current and lagged significant effects range from 0.291 to 1.045. Moreover, we find that the coefficient estimates for the $INTER_{ij,t}$ dummies are positive (except for Wood, Paper, and Chemicals for some years) and, in many cases, statistically different from zero. We do not present the results for PQML with positive flows only, as these results were very close to those including zeros, as we established in Table 2. Also, in an online appendix we provide the results for PQML excluding observations for *ROW* and for interpolated output values, as these results also are very close to those in Table 4 (as we established earlier in Table 2).

[36] When exporting fixed costs are present and productivities are bounded from above (as, for example, in HMR), the gravity equations stated in Eqs. (1), (3), and (4) apply only for the positive trade flows. In Eq. (1), besides representing structural parameters $\chi_{ij,t}$ would also capture the corrections for country selection (the inverse Mills ratio) and firm heterogeneity, and those correction terms would have to be included as additional regressors in Eqs. (3) and (4).

[37] For this issue, it is critical to note that identification of the HMR terms $IMR_{ij,t}$, $Z_{ij,t}$, $Z_{ij,t}^2$, and $Z_{ij,t}^3$ relies on the variation of the export-status (positive or zero). The HMR approach was explored using our data set; however, convergence could not be achieved precisely because of the very small number of zeros.

Table 4
Estimates for sectoral manufactures trade flows.

4A. EIAs	(1) Food	(2) Textile	(3) Wood	(4) Paper	(5) Chemicals	(6) Minerals	(7) Metals	(8) Machinery
$EIA_{ij,t}$	0.366	0.569	−0.151	−0.109	0.099	−0.038	0.355	0.313
	(0.104)**	(0.108)**	(0.069)*	(0.048)*	(0.038)**	(0.070)	(0.048)**	(0.072)**
$EIA_{ij,t-4}$	0.253	0.459	0.203	0.121	0.221	0.390	0.202	0.351
	(0.048)**	(0.085)**	(0.039)**	(0.040)**	(0.051)**	(0.089)**	(0.049)**	(0.064)**
$EIA_{ij,t-8}$	0.202	0.323	0.010	0.332	0.187	0.145	0.280	0.297
	(0.034)**	(0.069)**	(0.081)	(0.068)**	(0.036)**	(0.026)**	(0.049)**	(0.057)**
$EIA_{ij,t-12}$	0.268	0.175	−0.152	−0.089	0.049	−0.011	0.067	0.098
	(0.028)**	(0.049)**	(0.050)**	(0.048)**	(0.063)	(0.055)	(0.042)	(0.041)*
$EIA_{ij,t-16}$	0.126	0.116	0.263	0.093	0.096	−0.062	0.095	0.087
	(0.047)**	(0.039)**	(0.054)**	(0.049)+	(0.055)+	(0.052)	(0.062)	(0.048)+
Total EIA	1.215	1.641	0.175	0.348	0.652	0.422	0.998	1.145
	(0.123)**	(0.207)**	(0.114)	(0.096)**	(0.065)**	(0.095)**	(0.109)**	(0.173)**
N	6692	6708	6596	6540	6724	6628	6564	6724
4B. EIAs, INTER, and other variables								
$EIA_{ij,t}$	0.320	0.349	−0.099	−0.160	0.021	−0.066	0.247	0.149
	(0.090)**	(0.067)**	(0.069)	(0.054)**	(0.047)	(0.080)	(0.049)**	(0.062)*
$EIA_{ij,t-4}$	0.205	0.324	0.256	0.131	0.255	0.368	0.109	0.290
	(0.032)**	(0.061)**	(0.038)**	(0.043)**	(0.047)**	(0.093)**	(0.050)*	(0.054)**
$EIA_{ij,t-8}$	0.151	0.130	0.138	0.297	0.279	0.181	0.214	0.210
	(0.036)**	(0.053)*	(0.058)*	(0.053)**	(0.046)**	(0.040)**	(0.055)**	(0.056)**
$EIA_{ij,t-12}$	0.202	−0.050	−0.032	−0.052	0.070	−0.036	−0.020	0.026
	(0.036)**	(0.058)	(0.059)	(0.046)	(0.047)	(0.056)	(0.063)	(0.052)
$EIA_{ij,t-16}$	0.167	0.083	0.246	0.075	0.093	−0.088	−0.007	−0.035
	(0.054)**	(0.047)+	(0.052)**	(0.062)	(0.065)	(0.070)	(0.063)	(0.044)
$INTER_{ij,1994}$	0.151	0.554	0.218	−0.024	−0.065	0.021	0.343	0.192
	(0.062)*	(0.118)**	(0.113)+	(0.064)	(0.068)	(0.102)	(0.088)**	(0.083)*
$INTER_{ij,1998}$	0.160	0.810	−0.106	0.071	−0.024	0.009	0.101	0.347
	(0.075)*	(0.161)**	(0.131)	(0.109)	(0.084)	(0.165)	(0.184)	(0.116)**
$INTER_{ij,2002}$	0.021	0.742	−0.184	0.034	−0.179	0.017	0.517	0.585
	(0.087)	(0.216)**	(0.169)	(0.163)	(0.119)	(0.284)	(0.106)**	(0.178)**
$DIST_{ij,1994}$	−0.039	−0.100	−0.101	0.041	0.046	−0.029	−0.092	−0.021
	(0.029)	(0.052)+	(0.048)*	(0.031)	(0.027)+	(0.046)	(0.047)*	(0.034)
$DIST_{ij,1998}$	−0.027	−0.106	0.063	0.035	0.107	0.046	0.137	0.028
	(0.035)	(0.069)	(0.061)	(0.046)	(0.034)**	(0.065)	(0.103)	(0.047)
$DIST_{ij,2002}$	0.021	−0.081	0.091	0.041	0.161	0.046	−0.093	−0.027
	(0.043)	(0.087)	(0.079)	(0.068)	(0.047)**	(0.104)	(0.062)	(0.069)
$LANG_{ij,1994}$	0.074	0.082	0.020	0.043	0.064	0.110	0.134	−0.021
	(0.049)	(0.096)	(0.097)	(0.049)	(0.062)	(0.115)	(0.081)+	(0.061)
$LANG_{ij,1998}$	0.153	0.089	−0.033	0.064	0.026	0.073	0.054	0.006
	(0.063)*	(0.153)	(0.109)	(0.080)	(0.066)	(0.109)	(0.108)	(0.083)
$LANG_{ij,2002}$	0.240	0.134	−0.055	0.053	0.213	0.076	−0.092	−0.029
	(0.060)**	(0.147)	(0.124)	(0.089)	(0.147)	(0.106)	(0.116)	(0.089)
$CNTG_{ij,1994}$	−0.053	−0.142	−0.214	−0.026	−0.029	−0.113	−0.232	−0.153
	(0.055)	(0.091)	(0.103)*	(0.056)	(0.069)	(0.083)	(0.071)**	(0.076)*
$CNTG_{ij,1998}$	−0.025	0.027	−0.108	−0.048	−0.024	−0.077	0.041	−0.133
	(0.067)	(0.151)	(0.128)	(0.099)	(0.082)	(0.112)	(0.127)	(0.111)
$CNTG_{ij,2002}$	0.049	0.045	−0.136	−0.112	−0.273	−0.149	−0.106	−0.341
	(0.067)	(0.161)	(0.148)	(0.128)	(0.138)*	(0.182)	(0.091)	(0.136)*
$CLNY_{ij,1994}$	−0.072	−0.048	−0.141	−0.073	−0.054	−0.058	−0.029	0.002
	(0.047)	(0.161)	(0.120)	(0.061)	(0.051)	(0.081)	(0.102)	(0.039)
$CLNY_{ij,1998}$	−0.051	−0.145	0.008	−0.105	−0.017	−0.058	0.007	−0.052
	(0.085)	(0.227)	(0.144)	(0.092)	(0.068)	(0.124)	(0.136)	(0.071)
$CLNY_{ij,2002}$	−0.087	−0.267	0.012	−0.045	−0.014	−0.148	0.076	−0.079
	(0.076)	(0.249)	(0.147)	(0.102)	(0.141)	(0.115)	(0.161)	(0.115)
Total EIA	1.045	0.836	0.509	0.291	0.718	0.358	0.544	0.640
	(0.115)**	(0.160)**	(0.146)*	(0.139)*	(0.122)**	(0.191)**	(0.117)**	(0.146)**
N	6692	6708	6596	6540	6724	6628	6564	6724

Notes: This table reports panel gravity estimates for the eight 2-digit ISIC categories of manufactures, 1990–2002. All specifications are estimated with PQML, pair (ij), exporter-year (it), and importer-year (jt) fixed effects and allow for phasing-in of the EIA effects. Fixed effects estimates are not reported for brevity. Robust standard errors, clustered by country pair, are reported in parentheses. See text for further details.

$^{+}$ $p < 0.10$
* $p < .05$
** $p < .01$

In sum, the main PQML results shown in column (3) of Table 3 for total manufactures are robust to using disaggregate trade flows. These results suggest that – after accounting for likely declining bilateral variable and fixed trade costs using a novel set of time-varying international border dummies capturing relative international-to-intranational trade-flow changes – EIAs still have economically and statistically significant partial effects on trade flows, but that ignoring the international border dummies tends to bias upward EIA estimates. Moreover, the novel time-varying international border dummies reveal the average cost of international relative to intranational trade has declined for several sectors.

6. Empirical results for aggregate trade flows

One of the limitations of the manufactures data set employed is the short time series; we are able to explain the effect of declining bilateral trade costs over a period of only 12 years.[38] However, data is available for a longer time series for bilateral aggregate goods trade flows. The drawback of using aggregate goods trade flows is that the only available (imperfect) proxy of gross output, from which to construct intranational trade flows, is Gross Domestic Product (GDP) – a "value-added" (not gross output) measure. Nevertheless, we thought it worthwhile in this section to evaluate the sensitivity of the earlier results for total manufactures trade flows using a shorter time series to use of aggregate goods trade flows for a longer period, keeping in mind the shortcoming in our measure of intranational trade.

The aggregate goods trade flow data are from COMTRADE. We nearly doubled the time period relative to earlier; it is 1991–2011. We have also more than doubled the number of countries covered from 40 to 89; the 89 countries included represent 98 percent of world output.[39] The GDP data come from the latest edition of the Penn World Tables 8.0.[40]

Table 5 provides the results. In the first column of results using PQML, we show that as before the current and lagged EIA dummies have economically and statistically significant positive effects. The introduction in the next specification of the international border dummies (and other gravity covariates) causes as expected the coefficient estimates of $EIA_{ij,t}$ and its lags to decline. The total EIA partial effect in column (2) of 0.496 is smaller than the comparable estimate in column (3) of Table 3 of 0.626; however, EIA effects tend to be lower for aggregate goods trade than for manufactures trade. It is worth noting the pattern of coefficient estimates for the five border dummies. All of the *INTER* coefficient estimates are positive and statistically significant.

The third and fourth columns report the results using OLS instead. As found earlier, EIAs' coefficient estimates are smaller using OLS relative to PQML and international border effects are larger. Consequently, the results for manufactures trade using a smaller number of countries and years are essentially comparable to the results using aggregate goods trade flows for a larger number of countries and years.

7. The "distance-elasticity" puzzle

One of the potential key contributions of this paper is the introduction of a variable $INTER_{ij,t}$ to account for likely declining trends in bilateral fixed and variable trade costs that are potentially increasing international relative to intranational trade. Moreover, as seen in Tables 3, 4B and 5, the further introduction of typical gravity-equation variables that likely explain international trade relative to intranational trade – *DIST*, *LANG*, *CNTG*, and *CLNY* – add virtually no further explanatory power once *INTER* is included.

However, what happens once *INTER* is excluded? Our expectation is that many of the typical gravity-equation variables matter. Table 6 provides the results from excluding *INTER*. Not surprisingly, $DIST_{ij,t}$ – the logarithm of bilateral distance interacted with the year dummies – replaces the influence of $INTER_{ij,t}$, and as before the other variables play a limited role. Yet, these results raise the possibility of addressing another important issue. One of the well-known puzzles in the empirical international trade literature is the "distance-elasticity puzzle." This puzzle is that – in spite of likely falling bilateral fixed and variable trade costs – a time series of cross-sectional estimates of a properly-specified gravity equation yield *rising* distance elasticities (in absolute values). That is, international trade in such cross sections declines more in response to distance in recent years relative to earlier years, cf., Disdier and Head (2008).

While several researchers have made attempts to solve the puzzle, Yotov, 2012 addressed the issue by including observations for *intranational* trade, along with including on the right-hand-side a variable measuring intranational distances. Such intranational trade flows and distances have actually been a common feature of several border-effect studies, but had not yet permeated the distance-elasticity literature. Yotov (2012) "solved" the distance-elasticity puzzle by

[38] The effective constraint is available production data.

[39] The 89 countries (with their respective labels in parentheses) include: Angola (AGO), Argentina (ARG), Australia (AUS), Austria (AUT), Azerbaijan (AZE), Bangladesh (BGD), Belarus (BLR), Belgium (BEL), Brazil (BRA), Bulgaria (BGR), Canada (CAN), Chile (CHL), China (CHN), Colombia (COL), Croatia (HRV), Czech Republic (CZE), Cyprus (CYP), Denmark (DNK), Dominican Republic (DOM), Ecuador (ECU), Egypt (EGY), Estonia (EST), Ethiopia (ETH), Finland (FIN), France (FRA), Germany (DEU), Ghana (GHA), Greece (GRC), Guatemala (GTM), Hong Kong (HKG), Hungary (HUN), India (IND), Indonesia (IDN), Iran (IRN), Iraq (IRQ), Ireland (IRL), Israel (ISR), Italy (ITA), Japan (JPN), Kazakhstan (KAZ), Kenya (KEN), Korea, Republic of (KOR), Kuwait (KWT), Lebanon (LBN), Lithuania (LTU), Latvia (LVA), Luxembourg (LUX), Macedonia (MKD), Malaysia (MYS), Malta (MLT), Mexico (MEX), Morocco (MAR), Netherlands (NLD), New Zealand (NZL), Nigeria (NGA), Norway (NOR), Oman (OMN), Pakistan (PAK), Peru (PER), Philippines (PHL), Poland (POL), Portugal (PRT), Qatar (QAT), Romania (ROU), Russia (RUS), Saudi Arabia (SAU), Serbia (SRB), Singapore (SGP), Slovak Republic (SVK), Slovenia (SVN), South Africa (ZAF), Spain (ESP), Sri Lanka (LKA), Sudan (SDN), Sweden (SWE), Switzerland (CHE), Syria (SYR), Tanzania (TZA), Thailand (THA), Tunisia (TUN), Turkey (TUR), Turkmenistan (TKM), Ukraine (UKR), United Kingdom (GBR), United States (USA), Uzbekistan (UZB), Venezuela (VEN), Vietnam (VNM), Zimbabwe (ZWE).

[40] These series are now maintained by the Groningen Growth and Development Centre and reside at http://www.rug.nl/research/ggdc/data/penn-world-table

Table 5
Estimates for aggregate goods trade flows, 1991–2011.

	(1) PQML	(2) PQML	(3) OLS	(4) OLS
$EIA_{ij,t}$	0.511 (0.070)**	0.085 (0.058)	0.172 (0.039)**	0.158 (0.039)**
$EIA_{ij,t-4}$	0.351 (0.045)**	0.228 (0.035)**	0.132 (0.034)**	0.136 (0.035)**
$EIA_{ij,t-8}$	0.295 (0.044)**	0.186 (0.040)**	−0.016 (0.036)	−0.020 (0.037)
$EIA_{ij,t-12}$	0.276 (0.060)**	0.016 (0.050)	0.070 (0.037)+	0.075 (0.038)*
$EIA_{ij,t-16}$	0.279 (0.022)**	−0.019 (0.041)	0.007 (0.041)	0.007 (0.043)
$INTER_{ij,1995}$		0.176 (0.056)**		0.311 (0.161)+
$INTER_{ij,1999}$		0.192 (0.075)*		0.252 (0.164)
$INTER_{ij,2003}$		0.460 (0.106)**		0.549 (0.180)**
$INTER_{ij,2007}$		1.055 (0.114)**		1.165 (0.196)**
$INTER_{ij,2011}$		1.206 (0.126)**		1.478 (0.227)**
$DIST_{ij,1995}$		−0.002 (0.020)		−0.008 (0.035)
$DIST_{ij,1999}$		−0.024 (0.026)		0.031 (0.035)
$DIST_{ij,2003}$		−0.062 (0.035)+		0.002 (0.038)
$DIST_{ij,2007}$		−0.078 (0.040)+		0.033 (0.039)
$DIST_{ij,2011}$		−0.021 (0.044)		0.045 (0.041)
$LANG_{ij,1995}$		0.040 (0.044)		0.002 (0.080)
$LANG_{ij,1999}$		0.004 (0.072)		0.128 (0.085)
$LANG_{ij,2003}$		−0.040 (0.060)		0.070 (0.092)
$LANG_{ij,2007}$		−0.096 (0.078)		0.098 (0.091)
$LANG_{ij,2011}$		−0.213 (0.094)*		−0.018 (0.093)
$CNTG_{ij,1995}$		−0.069 (0.066)		0.197 (0.119)+
$CNTG_{ij,1999}$		−0.134 (0.107)		0.058 (0.112)
$CNTG_{ij,2003}$		−0.211 (0.093)*		0.128 (0.121)
$CNTG_{ij,2007}$		−0.239 (0.087)**		−0.005 (0.134)
$CNTG_{ij,2011}$		−0.184 (0.102)+		0.078 (0.125)
$CLNY_{ij,1995}$		−0.060 (0.050)		−0.011 (0.095)
$CLNY_{ij,1999}$		0.027 (0.058)		−0.126 (0.084)
$CLNY_{ij,2003}$		−0.073 (0.074)		−0.159 (0.099)
$CLNY_{ij,2007}$		−0.165 (0.092)+		−0.207 (0.103)*
$CLNY_{ij,2011}$		−0.201 (0.103)+		−0.178 (0.104)+
Total EIA	1.713 (0.124)**	0.496 (0.096)**	0.364 (0.064)**	0.356 (0.068)**
N	44,078	44,078	41,726	41,726

Notes: This table reports panel gravity estimates with aggregate data of the effects of all EIAs that entered into force during the period 1991–2011. Columns (1) and (2) use the PQML estimator, while columns (3) and (4) use the OLS estimator. All specifications allow for phasing-in of the EIA effects. All specifications are estimated with pair (*ij*), exporter-year (*it*), and importer-year (*jt*) fixed effects. Fixed effects estimate are not reported for brevity. Robust standard errors, clustered by country pair, are in parentheses.

+ $p < 0.10$
* $p < .05$
** $p < .01$

noting the importance of measuring international distances *relative to* intranational distances, as theoretical foundations for gravity equations actually suggest.

However, a shortcoming of Yotov (2012) is that – by using a time-series of cross-sections – the author does not control for unobserved bilateral heterogeneity and consequently the results suffer from omitted variables bias. Moreover, the results did not account for endogenous EIA effects. In order to estimate the distance elasticity, country-pair fixed effects cannot be used, as they would subsume the cross-country variation in bilateral distance. Consequently, researchers typically include bilateral distance and many other pair-specific variables to explain trade cross-sectionally. This leads to potentially biased results.

To obtain unbiased estimates, our approach uses a panel with pair fixed effects. The pair fixed effects capture the cross-sectional negative impact of bilateral distance on trade flows. By then introducing a set of year dummies interacted with bilateral distance – $DIST_{ij,1994}$, $DIST_{ij,1998}$, and $DIST_{ij,2002}$ – we capture the changing effects of bilateral distance on trade flows *relative to the initial year*. Another way to look at this variable is that it is a time-varying measure of the changing costs of international trade relative to intranational trade, but using a continuous measure rather than the earlier employed international border dummies. Note that the variables $DIST_{ij,t}$ in Table 6B have economically and statistically significant positive effects. Moreover, the size of the coefficient estimates tends to increase from 1994 to 2002. For total manufactures in column (9), the sizes of the positive coefficients increase monotonically.

What do these coefficient estimates suggest? The country-pair fixed effects (whose coefficient estimates are not shown due to the very large number) pickup that bilateral distance has a negative effect on trade. However, the *positive* and typically increasing over time coefficient estimates for $DIST_{ij,1994}$, $DIST_{ij,1998}$, and $DIST_{ij,2002}$ indicate that the negative effect of bilateral distance is *declining* over time. In the context of the discussion above and gravity Eq. (1), these results are consistent with the costs of international trade falling over time relative to intranational trade, likely attributable to decreasing – but unobservable – bilateral fixed and variable trade costs, that are increasing international relative to intranational trade. In fact, the coefficient estimate in panel 6B for $DIST_{ij,2002}$ of 0.141 for total manufactures trade suggests that the average effect of distance on international relative to intranational trade has diminished by 13 percent over 12 years ($[(1/e^{0.141})-1] \times 100$ percent), or 1.2 percent per year. A comparison of our results with those in Yotov (2012, Table 2) also reveals lower estimates using our approach. For instance, in Yotov (2012) the distance elasticity for textiles fell 57 percent over 10 years. By contrast, our results in Panel 6B imply that the distance elasticity for textiles fell only 20 percent ($[(1/e^{0.229})-1] \times 100$ percent) over a similar 12-year period, and this was the largest estimated distance-elasticity decrease. In fact, the smaller declines in the distance elasticities in our study relative to Yotov (2012) suggest that the estimates in the latter study were biased upward by ignoring the effects of EIAs.

This result is novel because it is generated allowing the country-pair fixed effects to subsume the level effect of distance on trade flows, and allows this effect to differ across country pairs in the initial year.[41] Only two previous studies have included bilateral country-pair fixed effects to address the distance-elasticity puzzle. Carrere et al. (2009) account for unobserved bilateral heterogeneity in their OLS estimates as well as linear approximations of the multilateral price terms. They find rising distance elasticities; however, they do not account for EIAs or intranational trade and distances. Bosquet and Boulhol (2015) could not find declining distance elasticities using PQML including bilateral fixed effects, but that is likely attributable to their exclusion of intranational trade and distances. However, unlike Carrere et al. (2009) and Bosquet and Boulhol (2015), we include as in Yotov (2012) intranational trade and distances; this feature is important to find declining distance effects on international trade, because we identify the effects of the (change in) the distance elasticity using variation in international and intranational trade flows.

Finally, it is useful to compare the results in, say, the last column in Table 6B (for Total Manufactures) with those in column (3) of Table 3 (for Total Manufactures). The total EIA (partial) effect from Table 6B is 0.728, whereas the total EIA effect from Table 3 is 0.626; these differ by 0.104 and the difference is not statistically significant. Moreover, when we ignore the statistically insignificant EIA coefficient estimates, the total EIA effects differ by only 0.063 (0.648–0.585), which is also statistically insignificant. The importance of this is that the EIA coefficient estimates are apparently not biased much by the choice of specification. $DIST_{ij,t}$ and $INTER_{ij,t}$ are – as noted earlier – “two sides of the same coin.” Consequently, it is reasonable to conclude that the $DIST_{ij,t}$ coefficient estimates are also unbiased estimates of the declining effect of distance on international relative to intranational trade.

8. General equilibrium comparative statics

The paper so far has focused solely on partial, or direct, EIA effects. However, as Anderson and van Wincoop (2003) and (Head and Mayer, 2000) suggest, general equilibrium effects are also important for quantitative evaluation of trade-policy changes. Consequently, it is potentially interesting to determine how much the general equilibrium (GE) trade effects of EIAs are influenced by the different partial effect estimates with *and* without accounting for the declining effect of international borders on trade. In this section, we consider two quantitative general equilibrium (GE) comparative static exercises. First, we provide the GE trade effects of all observed EIAs that formed between 1990 and 2002 using the partial effect coefficient estimates ignoring $INTER_{ij,t}$ (and the other gravity covariates); these estimates are from column (1) of Table 3. We provide these GE effect estimates for all (bilateral) trade flows of each of our 40 countries and the *ROW* aggregate, the trade flows

[41] Coe et al. (2007) found evidence of declining distance elasticities by employing non-linear estimation (including PQML), although the declines ended between 1990 and 2000, but could not find declining distance elasticities using OLS. Berthelon and Freund (2008) found rising distance elasticities since 1985 using OLS. Larch etal (2015) found a declining distance effect using non-linear estimators, but not using OLS. However, Coe et al. (2007), Berthelon and Freund (2008), and Larch etal (2015) did not account for unobserved bilateral heterogeneity or for the endogeneity of EIAs, and did not include intranational trade flows and distances.

Table 6
Estimates to address the distance-elasticity puzzle.

6A. EIAs	(1) Food	(2) Textile	(3) Wood	(4) Paper	(5) Chemicals	(6) Minerals	(7) Metals	(8) Machinery	(9) Total
$EIA_{ij,t}$	0.366	0.569	−0.151	−0.109	0.099	−0.038	0.355	0.313	0.241
	(0.104)**	(0.108)**	(0.069)*	(0.048)*	(0.038)**	(0.070)	(0.048)**	(0.072)**	(0.038)**
$EIA_{ij,t-4}$	0.253	0.459	0.203	0.121	0.221	0.390	0.202	0.351	0.294
	(0.048)**	(0.085)**	(0.039)**	(0.040)**	(0.051)**	(0.089)**	(0.049)**	(0.064)**	(0.050)**
$EIA_{ij,t-8}$	0.202	0.323	0.010	0.332	0.187	0.145	0.280	0.297	0.248
	(0.034)**	(0.069)**	(0.081)	(0.068)**	(0.036)**	(0.026)**	(0.049)**	(0.057)**	(0.033)**
$EIA_{ij,t-12}$	0.268	0.175	−0.152	−0.089	0.049	−0.011	0.067	0.098	0.069
	(0.028)**	(0.049)**	(0.050)**	(0.048)+	(0.063)	(0.055)	(0.042)	(0.041)*	(0.045)
$EIA_{ij,t-16}$	0.126	0.116	0.263	0.093	0.096	−0.062	0.095	0.087	0.093
	(0.047)**	(0.039)**	(0.054)**	(0.049)+	(0.055)+	(0.052)	(0.062)	(0.048)+	(0.030)**
Total EIA	1.215	1.641	0.175	0.348	0.652	0.422	0.998	1.145	0.944
	(0.123)**	(0.207)**	(0.114)	(0.096)**	(0.065)**	(0.095)**	(0.109)**	(0.173)**	(0.085)**
N	6692	6708	6596	6540	6724	6628	6564	6724	6724
6B. EIAs and other variables									
$EIA_{ij,t}$	0.329	0.409	−0.122	−0.151	0.024	−0.066	0.242	0.181	0.133
	(0.089)**	(0.061)**	(0.074)+	(0.052)**	(0.046)	(0.077)	(0.051)**	(0.058)**	(0.042)**
$EIA_{ij,t-4}$	0.201	0.356	0.238	0.131	0.232	0.369	0.171	0.356	0.268
	(0.035)**	(0.059)**	(0.041)**	(0.038)**	(0.047)**	(0.087)**	(0.045)**	(0.044)**	(0.039)**
$EIA_{ij,t-8}$	0.158	0.193	0.106	0.309	0.274	0.182	0.230	0.262	0.247
	(0.035)**	(0.049)**	(0.061)+	(0.055)**	(0.045)**	(0.035)**	(0.052)**	(0.045)**	(0.030)**
$EIA_{ij,t-12}$	0.191	0.005	−0.071	−0.050	0.050	−0.036	0.031	0.095	0.047
	(0.037)**	(0.055)	(0.056)	(0.042)	(0.046)	(0.044)	(0.056)	(0.048)*	(0.034)
$EIA_{ij,t-16}$	0.160	0.091	0.233	0.075	0.078	−0.087	0.031	0.007	0.033
	(0.051)**	(0.040)*	(0.053)**	(0.062)	(0.065)	(0.062)	(0.057)	(0.041)	(0.031)
$DIST_{ij,1994}$	0.029	0.140	−0.004	0.030	0.018	−0.020	0.053	0.056	0.047
	(0.009)**	(0.027)**	(0.016)	(0.013)*	(0.009)*	(0.013)	(0.027)+	(0.014)**	(0.010)**
$DIST_{ij,1998}$	0.046	0.239	0.019	0.066	0.097	0.050	0.182	0.165	0.130
	(0.012)**	(0.039)**	(0.025)	(0.016)**	(0.013)**	(0.019)**	(0.041)**	(0.023)**	(0.013)**
$DIST_{ij,2002}$	0.031	0.229	0.015	0.056	0.087	0.053	0.121	0.203	0.141
	(0.015)*	(0.036)**	(0.033)	(0.026)*	(0.020)**	(0.028)+	(0.038)**	(0.027)**	(0.017)**
$LANG_{ij,1994}$	0.079	0.073	0.041	0.043	0.069	0.110	0.123	−0.032	0.032
	(0.052)	(0.088)	(0.103)	(0.049)	(0.062)	(0.116)	(0.077)	(0.059)	(0.047)
$LANG_{ij,1998}$	0.156	0.076	−0.037	0.067	0.028	0.073	0.045	−0.004	0.005
	(0.064)*	(0.144)	(0.111)	(0.080)	(0.067)	(0.108)	(0.107)	(0.075)	(0.070)
$LANG_{ij,2002}$	0.241	0.122	−0.059	0.054	0.218	0.077	−0.099	−0.047	0.024
	(0.060)**	(0.138)	(0.126)	(0.088)	(0.149)	(0.106)	(0.109)	(0.082)	(0.071)
$CNTG_{ij,1994}$	0.040	0.177	−0.077	−0.044	−0.069	−0.100	−0.021	−0.040	−0.039
	(0.040)	(0.049)**	(0.086)	(0.037)	(0.055)	(0.069)	(0.045)	(0.053)	(0.043)
$CNTG_{ij,1998}$	0.076	0.483	−0.158	−0.005	−0.032	−0.071	0.082	0.057	0.056
	(0.050)	(0.124)**	(0.110)	(0.073)	(0.067)	(0.062)	(0.060)	(0.079)	(0.072)
$CNTG_{ij,2002}$	0.061	0.440	−0.222	−0.096	−0.377	−0.139	0.185	−0.018	−0.060
	(0.047)	(0.108)**	(0.127)+	(0.085)	(0.125)**	(0.083)+	(0.062)**	(0.079)	(0.068)
$CLNY_{ij,1994}$	−0.083	0.012	−0.157	−0.071	−0.059	−0.057	−0.007	0.019	−0.025
	(0.048)+	(0.139)	(0.123)	(0.061)	(0.051)	(0.079)	(0.090)	(0.033)	(0.037)
$CLNY_{ij,1998}$	−0.062	−0.071	0.012	−0.111	−0.021	−0.057	0.016	−0.018	−0.017
	(0.086)	(0.189)	(0.146)	(0.093)	(0.068)	(0.122)	(0.132)	(0.065)	(0.061)
$CLNY_{ij,2002}$	−0.088	−0.198	0.015	−0.047	−0.024	−0.147	0.116	−0.021	−0.038
	(0.077)	(0.211)	(0.151)	(0.102)	(0.141)	(0.110)	(0.142)	(0.112)	(0.060)
Total EIA	1.039	1.052	0.383	0.313	0.657	0.362	0.705	0.901	0.728
	(0.113)**	(0.124)**	(0.143)*	(0.117)*	(0.110)**	(0.131)**	(0.109)**	(0.102)**	(0.085)**
N	6692	6708	6596	6540	6724	6628	6564	6724	6724

Notes: This table reproduces the results from Table 5 but without the *INTER* variable. All specifications are estimated with PQML, pair (ij), exporter-year (it), and importer-year (jt) fixed effects and allow for phasing-in of the EIA effects. Fixed effects estimates are not reported for brevity. Robust standard errors, clustered by country pair, are reported in parentheses. See text for further details.

+ $p < 0.10$

* $p < .05$

** $p < .01$

among all countries in the world (denoted World), the trade flows among countries where EIAs formed, and the trade flows among countries where EIAs did *not* form. In the spirit of Head and Mayer (2000), we provide three GE comparative static estimates: a “Partial Trade Impact” (PTI) estimate that ignores (endogenous) changes in multilateral resistance (MR) terms and GDPs, a “Multilateral Trade Impact” (MTI) estimate that accounts for changes in multilateral resistance terms but

Table 7
General equilibrium trade effects of globalization.

	EIAs (1) PTI	(2) MTI	(3) GETI	EIAs and Borders (4) PTI	(5) MTI	(6) GETI
ARG	30.286	25.718	26.173	83.446	65.587	67.330
AUS	0.000	−0.962	−0.859	64.078	47.315	48.924
AUT	55.157	32.435	34.741	83.461	42.766	46.537
BGR	62.978	54.326	55.159	110.020	85.872	88.154
BLX	5.131	1.524	1.872	41.793	13.331	16.084
BOL	45.598	32.906	34.170	95.865	62.975	66.076
BRA	13.280	10.444	10.736	72.969	57.948	59.437
CAN	56.880	42.363	43.795	97.162	66.135	69.073
CHE	2.973	0.638	0.851	41.867	23.980	25.628
CHL	17.045	12.637	13.081	80.715	57.715	59.899
CHN	0.000	−0.847	−0.753	60.512	43.799	45.407
COL	6.427	3.611	3.904	65.115	49.207	50.770
CRI	1.878	−0.040	0.126	62.528	34.251	37.011
DEU	11.504	7.138	7.563	53.179	32.038	33.997
DNK	12.674	6.041	6.694	59.460	26.979	29.982
ECU	27.982	21.954	22.554	87.145	63.838	66.060
ESP	41.676	32.248	33.164	84.665	58.209	60.650
FIN	49.619	34.853	36.289	98.718	60.719	64.155
FRA	7.823	4.741	5.035	45.201	27.596	29.228
GBR	6.480	3.690	3.956	61.102	38.218	40.346
GRC	20.390	14.377	14.969	69.085	43.146	45.535
HUN	92.769	55.076	58.712	148.591	77.108	83.434
IRL	3.135	0.826	1.037	48.466	17.363	20.413
ISL	2.624	0.425	0.620	60.613	35.755	38.025
ISR	30.825	22.766	23.563	87.760	57.427	60.229
ITA	8.736	5.379	5.702	57.376	36.998	38.886
JPN	0.000	−1.241	−1.106	60.981	48.016	49.308
KOR	0.000	−0.918	−0.820	58.040	41.402	43.005
MAR	22.205	15.624	16.272	72.223	46.202	48.603
MEX	95.290	45.032	50.115	120.119	50.214	56.916
NLD	5.694	1.905	2.274	46.404	16.215	19.119
NOR	3.957	0.256	0.584	53.405	28.797	31.013
POL	98.102	72.814	75.198	132.552	87.433	91.486
PRT	45.379	31.282	32.659	91.281	55.687	58.930
ROM	62.123	53.850	54.654	116.898	92.058	94.416
ROW	0.000	−0.935	−0.847	62.238	39.466	41.638
SWE	45.178	27.937	29.681	95.432	50.403	54.502
TUN	59.379	41.008	42.821	112.449	68.369	72.375
TUR	53.335	42.794	43.807	111.234	79.769	82.659
URY	51.072	38.388	39.636	95.920	65.776	68.614
USA	19.055	13.605	14.190	74.464	53.666	55.743
World	14.365	9.583	10.079	63.852	40.389	42.650
EIA	73.174	53.587	55.603	108.855	68.449	72.295
No EIA	0.000	−2.262	−2.044	51.413	32.006	33.868

Notes: This table reports the general equilibrium trade effects of introducing all EIAs that entered into force during the period of investigation from 1990 to 2002. Manufacturing output values are taken from 1990. The first scenario, labeled "EIAs," is based on estimates from column (1) in Table 3. The second scenario, labeled "EIAs and Borders" is based on estimates from column (3) in Table 3. For both scenarios, we report three different trade impacts: the Partial Trade Impact (PTI), which only takes the partial, direct effect into account; the Multilateral Trade Impact (MTI), which takes changes in the multilateral resistances (MR) into account, but holds GDPs constant; and the General Equilibrium Trade Impact (GETI), where MRs and GDPs adjust. The row "World" gives results for the average change in total world trade flows, the row "EIA" reports the average change in total trade flows where EIAs are formed, and the row "No EIA" gives the average change in total trade flows where no EIAs are formed.

ignores changes in GDPs, and a "General Equilibrium Trade Impact" (GETI) estimate that accounts for changes in MR terms and GDPs. Second, we provide the GE trade effects allowing for $INTER_{ij,t}$; these are based on coefficient estimates from column (3) of Table 3. Hence, the latter estimates account for GE effects of both formations of EIAs and the declining effects of international borders. We then compare the results.

Table 7 provides the results from the two comparative static exercises. Vertically, the results are reported first for the trade flows of each of the 40 countries and *ROW* aggregate, then the trade flows of the world, then the trade flows among countries where EIAs formed, and then the trade flows among countries where EIAs did not form. Columns (1)–(3) report the results for each of PTI, MTI, and GETI for the first exercise, ignoring the effects of $INTER_{ij,t}$. Columns (4)–(6) report the results for the second exercise, formations of EIAs and declining effects of borders (denoted "EIAs and Borders"). We discuss first the results in columns (1)–(3).

Several results from columns (1) to (3). First, the PTIs differ across countries even though the "partial effect" (or EIA coefficient estimate) is the same. The reason is that the PTI for each country reflects the total trade (partial) effect for that country. Countries that formed more agreements (like Hungary, Mexico, and Poland) had larger total trade changes, and thus larger PTIs. For the world on average, trade increased only about 14 percent, due to the fact that EIAs are rare events; note that the PTI for *ROW* was 0. Trade increases about 73 percent among countries where EIAs formed. For trade among countries where no EIAs formed, the PTI was 0.

Second, as expected, the MTIs are smaller than the PTIs. As the countries' multilateral resistance terms decline with the EIAs, the PTIs are offset, cf., Head and Mayer (2014). Typically for our comparative statics, the MR effects reduce the trade impacts by about one-third. For trade among countries where EIAs did not form, MTIs expectedly fell.

Third, the positive GDP effects from the EIAs did raise trade. But as found in previous like estimates, such effects are quantitatively small. Hence, the GETIs are quite similar to the MTIs.

We now consider a comparison of the trade effects from just EIAs versus those from EIAs *alongside* declining effects of international borders. First, as expected the PTIs are larger in columns (4)–(6) vis-á-vis their respective counterparts. These PTIs reflect both EIA effects as well as falling international border effects.

Second, note that, for trade among countries where EIAs did not form, the trade effect of declining borders is about 51 percent. Comparing this PTI with that for trade among countries where EIAs formed, the marginal trade gain from having EIAs is only about 57 percent (108.9–51.4). This reflects that the PTI for trade among countries where EIAs formed is less in the second exercise, as expected; the PTI from EIAs is smaller once declining border effects are accounted for, consistent with our earlier empirical findings.

Third, as before, the MTIs are smaller than the PTIs in columns (4)–(6) as expected. The decline is again about one-third. Also, the GETIs are slightly larger than the MTIs in columns (4)–(6), as expected.

In summary, the basic insights from the earlier analysis hold up in a GE comparative static analysis. Not surprisingly, the differential effect across countries is largely driven by differences in the PTIs, which by construction capture the extent of EIA liberalizations that countries pursued during the 1990–2002 period. These results support the notion that consistent and unbiased estimates of partial EIA effects are important for subsequent GE analysis. Moreover, the GETI for the world accounting for EIAs and the declining effect of international borders of 43 percent, compared to only 10 percent for EIAs alone, suggests that the declining GE effects of distance are non-trivial.

9. Conclusions

Using a common gravity-equation specification, we have attempted to provide consistent and precise estimates of the (partial) effects of three important factors in international trade that typically have been addressed in three somewhat separate literatures. First, we have improved upon the specification in BB for estimating the effects of EIAs on international trade flows by controlling now for unobservable exogenous *time-varying* bilateral fixed and variable export costs that may have increased international relative to intranational trade; our results suggest that previous estimates of EIAs' effects were biased upward. Using our econometrically preferred estimator (PQML), the partial effect of an EIA falls by 30 percent; using OLS (as in BB and BBF), the partial EIA effect falls by much less.

Second, our novel approach allows us to estimate precisely the declining effect of international borders on international relative to intranational trade, allowing for unobserved bilateral country-pair heterogeneity and endogenous EIAs. While previous authors have found evidence of declining border effects (also, as typical, using data sets including intranational along with international trade flows), one of the shortcomings of these previous studies is omitted variables bias due to not accounting for differences in initial border-effect levels nor for endogenous EIAs. Our results suggest that previous estimates of the declining effect of international borders were biased upward, and we find the depressing effects of international borders on international trade have declined by 2.4 percent per year from 1990 to 2002.

Third, in an extensive sensitivity analysis, we introduce another method for accounting for unobserved time-varying declines in the costs of bilateral international relative to intranational trade. Accounting for endogenous EIAs and unobserved country-pair heterogeneity, we provide plausible estimates of the declining effect of distance on international trade, providing empirical support for the elusive declining "distance elasticity" of international trade. While our approach recognizes as in Yotov (2012) the importance of including intranational trade flows and using PQML in estimation in order to find declining distance elasticities, our novel contribution here is to account for unobserved country-pair heterogeneity and endogenous EIAs. We find that the estimates in Yotov (2012) of the declining effects of distance on international relative to intranational trade have been biased upward by not accounting for endogenous EIAs and unobserved bilateral heterogeneity. Our results suggest that the effect of distance on international trade has declined by 1.2 percent annually.

Just as BB contributed to the literature by emphasizing the importance of accounting for exporter-year, importer-year, and country-pair fixed effects in estimating the (partial) effects of EIAs, our hope is that – going forward – subsequent analyses account for all of the following using panel techniques:

1. Exporter-year and importer-year fixed effects to account for endogenous prices and unobserved time-varying exporter and importer multilateral heterogeneity;

2. Country-pair fixed effects or country-pair fixed effects interacted with a time trend to account for unobserved time-invariant or time-varying, respectively, bilateral effects (that subsume pair-specific border effect levels);
3. Intranational as well as international trade flows and bilateral distances, so that international border dummies can be introduced to account for declining international *relative to* intranational bilateral trade costs other than EIAs; and
4. PQML estimation to account for heteroskedasticity bias (owing to Jensen's inequality) and to account for zero trade flows.

Finally, it is important to note that estimates of general equilibrium effects of trade-cost changes are very sensitive to partial effect estimates. Anderson and van Wincoop (2003) and Bergstrand et al. (2013) provide extensive analyses of the sensitivity of general equilibrium estimates of trade-flow changes to underlying partial effects. Our final robustness analysis confirmed that general equilibrium estimates of EIA formations are also sensitive to accounting properly for declining effects of international borders.

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Appendix A. Supplementary data

Supplementary data associated with this article can be found in the online version at http://dx.doi.org/10.1016/j.euroecorev.2015.06.003.

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