



Economic integration agreements and the margins of international trade



Scott L. Baier^{a,b}, Jeffrey H. Bergstrand^{c,d,e,*}, Michael Feng^a

^a John E. Walker Department of Economics, Clemson University, Clemson, SC 29634, USA

^b Federal Reserve Bank of Atlanta, Atlanta, GA, USA

^c Department of Finance, Kellogg Institute for International Studies, University of Notre Dame, Notre Dame, IN 46556, USA

^d Department of Economics, Kellogg Institute for International Studies, University of Notre Dame, Notre Dame, IN 46556, USA

^e CESifo, Munich, Germany

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ABSTRACT

One of the main policy sources of trade-cost changes is the formation of an economic integration agreement (EIA), which potentially affects an importing country's welfare. This paper: (i) provides the first evidence using gravity equations of *both* intensive and extensive (goods) margins being affected by EIAs employing a panel data set with a large number of country pairs, product categories, and EIAs from 1962 to 2000; (ii) provides the first evidence of the differential (partial) effects of various "types" of EIAs on these intensive and extensive margins of trade; and (iii) finds a novel differential "timing" of the two margins' (partial) effects with intensive-margin effects occurring sooner than extensive-margin effects, consistent with recent theoretical predictions. The results are robust to correcting for potential sample-selection, firm-heterogeneity, and reverse causality biases.

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1. Introduction

The gravity equation has long dominated the international trade literature as the main econometric approach toward estimating ex post the "partial" (or direct) effects of economic integration agreements and other natural and policy-based bilateral trade costs on aggregate bilateral trade flows.¹ Economic integration agreements (EIAs) refer broadly to preferential trade agreements, free trade agreements, customs unions, common markets, and economic unions.² Recently, Baier and Bergstrand (2007) demonstrated that estimation (ex post) of the (partial) effects of EIAs suffered from endogeneity bias, mainly due to self-selection of country-pairs' governments into agreements. They showed that – after accounting for such bias using panel techniques – EIAs had much larger effects on trade flows than revealed in the earlier

gravity equation literature and these estimates were more precise. Anderson and Yotov (2011) confirmed these findings using panel data also. Such results followed in the footsteps of empirical trade studies such as Trefler (1993) and Lee and Swagel (1997) that showed that previous estimates of trade-policy liberalizations on imports were underestimated considerably due to endogeneity bias.

While such positive estimates for EIA dummy variables were interpreted in the context of either Armington or Krugman models as EIAs increasing trade volumes of existing homogeneous firms (i.e., the "intensive margin"), consideration of zeros in bilateral trade, fixed export costs, and firm heterogeneity have led researchers more recently to examine various "extensive margins" of trade. Such extensive margins fall under three general categories: country, goods (or products), and firm. The existence of zeros in aggregate bilateral trade flows among many country pairs has led some researchers to explore the probability that a pair of countries trades at all; to the extent that an EIA affects this probability, this changes the *country* extensive margin of trade and potentially economic welfare.

A second margin is known as the "goods" margin of trade. Hummels and Klenow (2005), or HK, introduced this notion by examining zeros in bilateral trade flows at *highly disaggregated product-category levels*. The motivation for HK was to explore in a cross section of a large number of products and among a large number of U.S. trading partners a fundamental question: Do large economies export more because they export

* Corresponding author at: Department of Finance, Mendoza College of Business, University of Notre Dame, Notre Dame, IN 46556, USA.

E-mail addresses: sbaier@clemson.edu (S.L. Baier), bergstrand.1@nd.edu (J.H. Bergstrand), fdayu@clemson.edu (M. Feng).

¹ Partial (or direct) effects refer to the absence of general-equilibrium (or indirect) effects; see Anderson and van Wincoop (2003) and Baier and Bergstrand (2009) on partial versus general equilibrium trade effects of trade-cost changes. "Aggregate" refers to all "goods" (or industries or product categories).

² In this study, we use the term "preferential trade agreement" to denote one with only partial liberalization (not free trade).

larger quantities of a given good (i.e., intensive goods margin) or a wider set of goods (extensive goods margin)?³ They found in their cross section that about 60% of larger exports of large economies was attributable to the extensive goods margin; specifically, as the exporter country's economic size grew, it exported a larger number of product categories (or "goods") to more markets. However, HK did not investigate the relationship between trade liberalizations and the intensive and extensive goods margins of trade. The purpose of this paper is to address this shortcoming.

In this paper, we explore the impact of EIAs on aggregate trade flows, intensive (goods) margins, and extensive (goods) margins for a large number of goods, country pairs, and years.⁴ This is important for at least three reasons. First, the relative impacts on intensive versus extensive margins of trade liberalizations may matter for estimating the welfare gains from trade. Traditionally, the welfare gains from trade liberalizations in models such as Armington and Krugman arise due to terms-of-trade changes; this is summarized succinctly in [Arkolakis et al. \(2012\)](#). In [Eaton and Kortum \(2002\)](#), trade liberalizations increase welfare due to an increase in economic efficiency ala the Dornbusch–Fisher–Samuelson model. In the [Melitz \(2003\)](#) model, trade liberalizations lead to gains due to firm heterogeneity and resulting increases in aggregate productivity. Second, while [Arkolakis et al. \(2012\)](#) recently argued that the welfare gains are iso-morphic across many modern quantitative trade models, they note that the gains can vary across models allowing heterogeneous firms depending upon the type of Melitz model; hence, the distinction between intensive margin effects and extensive margin effects is important for ultimately quantifying with more precision the "gains from trade."⁵ Third, the HK analysis limited itself to a cross section. In a panel, however, intensive margin and extensive margin effects of EIAs may have differential "timings." For instance, [Arkolakis et al. \(2012\)](#) recently introduced staggered "Calvo pricing" into their Ricardian model of trade and showed that the intensive margin likely reacts sooner to trade liberalizations than does the extensive margin. Moreover, since the two margins have different "trade elasticities," the quantitative path of the welfare gains is time sensitive.

Our paper extends the literature by offering three potential empirical contributions. First, we extend the [Baier and Bergstrand \(2007\)](#) panel econometric methodology for the (partial) effects of EIAs on aggregate trade flows using a gravity equation to examine in a setting with a large number of country pairs the effects of virtually all EIAs on the extensive and intensive goods margins, using the HK trade-margin-decomposition methodology. In the context of an econometric analysis, we are the first to find economically and statistically significant EIA effects on both the intensive and extensive (goods) margins in the context of a large number of country pairs, EIAs, and years.

Second, we examine the effects of various types of EIAs – one-way preferential trade agreements (OWPTAs), two-way preferential trade agreements (TWPTAs), free trade agreements (FTAs), and a variable for customs unions, common markets and economic unions (CUCMECUs) – on trade flows, extensive margins, and intensive margins.⁶ While two recent studies have adapted the Baier–Bergstrand methodology for estimating the effect of differing "types" of EIAs on bilateral aggregate trade flows, no econometric study has examined the effect of various types of EIAs on the (goods) extensive and intensive margins of trade using a

³ Each "good" was a 6-digit SITC category. They also explored the effects of country size and per capita GDP on the quality of goods exported, as well as the two margins.

⁴ Because firm-level data is not available for a large number of country-pairs for a large number of years, we are constrained to investigating EIAs impacts on products defined at the 4-digit SITC category level, as in [Hillberry and McDaniel \(2002\)](#), [Kehoe and Ruhl \(2009\)](#), and [Foster et al. \(2011\)](#) discussed below.

⁵ For instance, welfare estimates could be sensitive to the presence or absence of intermediates or multiple sectors. See also [Melitz and Redding \(2013\)](#) and [Feenstra and Weinstein \(2013\)](#).

⁶ The HK methodology is based on [Feenstra \(1994\)](#). Due to few observations on common markets and economic unions, we combine these two types of "deeper" EIAs with customs unions to form the variable CUCMECU, representing "deep" EIAs.

large number of country pairs and EIAs.⁷ Neither [Helpman et al. \(2008\)](#) nor [Egger et al. \(2011\)](#) distinguished among various types of EIAs in their analyses of country intensive and extensive margins. We find not only that deeper EIAs have larger trade effects than FTAs, and the latter have larger effects than (partial) two-way and one-way PTAs, but we distinguish between these various trade effects at the extensive and intensive margins using a panel of (disaggregate) bilateral trade flows from 1962 to 2000 covering 98% of world exports.

Third, [Bernard et al. \(2009\)](#) is likely the only empirical study to date to explore the "timing" of extensive and intensive margin responses to shocks. Using cross-sectional variation to examine long-run aspects, [Bernard et al. \(2009\)](#) find that variation in trade flows across country pairs is explained largely by the extensive margin, using firm-level data (the "firm" margin); this result is consistent with HK using their "goods" margin. But using *time-series* variation, [Bernard et al. \(2009\)](#) find that a larger proportion of trade variation can be explained by the intensive margin at short (five-year) time intervals. They show that, following the Asian financial crisis of 1997, virtually all of the variation in trade flows within 2–3 years could be explained by the *intensive* margin. This finding is consistent with two recent theoretical studies arguing that the low trade-cost elasticity found in macroeconomic analyses of business cycles should be associated with the intensive margin of trade compared with the relatively higher trade-cost elasticity found in international trade, which reflects the intensive and extensive margin effects.⁸ In this paper, we allow for differential "timing" of EIA effects using panel data. We find the first comprehensive empirical evidence that the shorter-term effects of EIAs on trade flows are more at the (goods) intensive margin and longer-term effects are more at the extensive margin (the latter entailing either fixed export costs or staggered "Calvo pricing" by consumers), consistent with intuition and results in [Bernard et al. \(2009\)](#). Moreover, our results shed empirical light on theoretical conjectures for the *relative* quantitative effects on intensive and extensive margins of variable trade cost changes in a Melitz-type model. Finally, we show our results are robust to potential country-selection, firm-heterogeneity, and reverse causality biases.

The remainder of this paper is as follows. [Section 2](#) discusses our methodology, based on the HK linear trade-margins-decomposition method and the [Baier and Bergstrand \(2007\)](#) approach for estimating partial effects of EIAs on trade flows in gravity frameworks. [Section 3](#) discusses data and measurement issues. [Section 4](#) provides the main empirical results and findings from three sensitivity analyses. [Section 5](#) concludes.

2. Methodology

Only three empirical studies have explored the effects of trade liberalizations – and, in particular, EIAs – on the intensive and extensive goods margins of trade using the HK methodology. The earliest study using the HK decomposition to explore this issue is [Hillberry and McDaniel \(2002\)](#), focusing solely on the North American Free Trade Agreement (NAFTA). Although they do not attempt to establish causal effects from NAFTA to trade increases, they provide a decomposition of post-NAFTA trade among the three partners into goods intensive and extensive margins using 4-digit Standard International Trade Classification (SITC) data. They find evidence of both margins changing between 1993 and 2001. [Kehoe and Ruhl \(2009\)](#) examined NAFTA, the earlier Canada–U.S. FTA trade liberalization, and some structural

⁷ The two studies that extended the Baier–Bergstrand framework to differing types of EIAs are [Magee \(2008\)](#) and [Roy \(2010\)](#); both found that customs unions had larger aggregate trade flow effects than FTAs. However, neither study examined extensive versus intensive margin issues.

⁸ [Ruhl \(2008\)](#) explains the delayed effect of the extensive-margin effects to fixed export costs on the supply side, while [Arkolakis et al. \(2011\)](#) explain the delayed effect of the extensive-margin effects to "Calvo pricing" by consumers on the demand side.

transformations using a modified version of the HK decomposition methodology and applied to a series of cross sections. Similar to Hillberry and McDaniel (2002), they do not conduct an econometric analysis trying to explain the effect of NAFTA (or the Canada–U.S. FTA) on trade flows conditional on other variables. They decompose actual goods extensive and intensive margin changes post-agreement also using 4-digit SITC data for goods categories from Feenstra et al. (2005). They find significant evidence of both extensive and intensive margin changes using their modified HK decomposition methodology. Both studies' evidence of goods intensive and extensive margins of trade expanding following the signing of NAFTA suggests the need for a comprehensive econometric analysis (conditional on other covariates) of the effects of EIAs in general on the goods intensive and extensive margins of trade, in the spirit of HK's original analysis of the effect of country size and per capita GDP on the two goods' margins.

The only study to our knowledge that like us uses a data set for a large number of country pairs and years, a large number of EIAs, and the HK methodology is Foster et al. (2011). However, the partial effect they found of an EIA on the goods extensive margin was an economically insignificant 10%, and they found virtually no effect of EIAs on the intensive margin. The latter result is a puzzle because it is existing exporters and importers that seek EIAs and the theoretical studies noted above suggest that the shorter-term effect of EIAs should be on the intensive margin. Yet, there are several differences between our study and theirs. First, their estimation was based upon a traditional gravity-equation specification ignoring recent theoretical developments that emphasize the importance of relative price or “multilateral resistance” terms. Our paper is based upon state-of-the-art gravity-equation specifications, such as discussed in Arkolakis et al. (2012). Second, Foster et al. (2011) use a short three-year window on both sides of the EIA formation, and consequently can only capture short-term EIA effects; this likely explains their economically small partial effects but does not explain finding only an extensive margin effect. As shown in Baier and Bergstrand (2007), EIAs can take 10–15 years to have their full impact on aggregate bilateral trade flows. Moreover, by allowing longer lags, we can distinguish between short-term vs. longer-term effects. Third, Foster et al. (2011) examine the impact of EIAs using a single dummy variable; we use multiple EIA variables to distinguish the effects of one-way PTAs, two-way PTAs, FTAs, and deeper EIAs on aggregate trade flows, extensive margins, and intensive margins. Even disregarding the potential endogeneity biases introduced by their ignoring relative prices, their study did not distinguish between various “types” of EIAs and did not distinguish between the “timing” of intensive and extensive margin effects. Indeed, Foster et al. (2011) suggest in their concluding paragraph that examination of the shorter-run versus longer-run effects and accounting for the differing “depth and breadth” of EIAs would be useful extensions.

2.1. The gravity equation

As Arkolakis et al. (2012) and others have noted, there is a number of models commonly considered in the trade literature (such as Armington, Krugman, Ricardian, Melitz) that yield iso-morphic gravity equations. Following Arkolakis et al. (2012), using a Melitz model one can generate a standard gravity equation:

$$X_{ijt}^m = N_{it}^m \gamma_{jt}^m \left(\frac{(a_{lit}^m)^{-\gamma^m} w_{it}^{-\gamma^m} \tau_{ijt}^{-\gamma^m} f_{ijt}^{-[\gamma^m/(\sigma^m-1)]}}{\sum_{k=1}^K N_{kt}^m (a_{lkt}^m)^{-\gamma^m} w_{kt}^{-\gamma^m} \tau_{kjt}^{-\gamma^m} f_{kjt}^{-[\gamma^m/(\sigma^m-1)]}} \right) \quad (1)$$

where X_{ijt}^m is the trade flow from i to j in year t in “good” m , N_{it}^m is the number of firms in i (exporting and non-exporting) that produce output in good m , Y_{jt}^m is the expenditure in j on good m , a_{lit}^m (defined as unit input requirements of labor) is the lower bound of the Pareto distribution of productivities in m in i , γ^m is an index of productivity heterogeneity among firms in good m , w_{it} is the wage rate in i , τ_{ijt} is variable

trade costs of exporting i 's products into j , f_{ijt} is fixed export costs from i to j , and σ^m is the elasticity of substitution in consumption.⁹ Note that the relative price term in large parentheses is a standard representation of relative prices in the gravity equation, but now also reflecting productivity heterogeneity (through a_{lit}^m and γ^m) and fixed exporting costs (f_{ijt}), cf., Melitz (2003), Chaney (2008), Redding (2011), and Arkolakis et al. (2012).

In the context of these models, variable trade costs, τ_{ijt} , affect X_{ijt}^m via both the intensive and extensive margins. As Chaney (2008) demonstrates in his Melitz-type model, $\gamma^m = (\sigma^m - 1) + [\gamma^m - (\sigma^m - 1)]$. $\sigma^m - 1$ represents the intensive margin elasticity of variable trade costs whereas $\gamma^m - (\sigma^m - 1)$ is the extensive margin elasticity of variable trade costs. For finite means in the theory, $\gamma^m/(\sigma^m - 1)$ must exceed 1. Empirically, Chaney (2008) notes empirical estimates of $\gamma^m/(\sigma^m - 1)$ range between 1.5 and 2. Hence, these models suggest that the intensive margin variable-trade-cost elasticity should be larger than the extensive margin variable-trade-cost elasticity. For instance, if $\gamma = 1.5(\sigma - 1)$, then the intensive margin elasticity is twice as large as the extensive margin elasticity; if $\gamma = 2(\sigma - 1)$, then the intensive margin elasticity is equal to the extensive margin elasticity.¹⁰

Interestingly, the theoretical result that the (variable-trade-cost) intensive margin elasticity should be at least as large as the extensive margin elasticity conflicts with the empirical results for the EIA partial effects in Foster et al. (2011) discussed earlier. We evaluate empirically this implication later, a potential contribution of this paper.

2.2. Accounting for endogenous EIAs: the Baier–Bergstrand methodology

Baier and Bergstrand (2007), or BB, re-evaluated usage of the gravity equation econometrically for estimating partial effects of EIAs on pairs of countries' trade flows.¹¹ The first of two main contributions was that self-selection of country-pairs into EIAs (cf., Baier and Bergstrand, 2004) likely created a significant endogeneity bias in previous gravity-equation estimates of the (partial) effects of EIAs on trade flows. This is precisely the concern raised in Arkolakis et al. (2012, section V) for gravity-equation estimates of trade elasticities; the observed variable trade cost measure may be correlated with unobservable trade costs hidden in the gravity equation's error term. The second main contribution of BB was that – given the slow-moving nature of EIAs' determinations – gravity equation estimation could use panel techniques and data to avoid endogeneity bias and also capture lagged influences, incorporating either bilateral fixed effects (in a log-levels specification) or first-differencing to account for time-invariant bilateral unobservable RHS variables, as well as incorporating exporter-time and importer-time effects to capture time-varying unobservable “multilateral price/resistance” terms of the exporter and importer. BB showed that EIAs on average increased two members' bilateral trade by approximately 100% after 10–15 years. Such a panel approach allows estimates of the “timing” of EIAs' effects on trade flows between short run and long

⁹ We refer here to Eq. (23) in Arkolakis et al. (2012), under the assumption that fixed export costs are paid in the importing country (i.e., $\mu = 0$).

¹⁰ However, f_{ijt} works entirely through the extensive margin, so that there is no clear theoretical hypothesis for the relative sizes of intensive and extensive margin effects of a given EIA formation. Yet, in light of estimates of $\gamma/(\sigma - 1)$ between 1.5 and 2, this additional fixed-trade-cost elasticity can range feasibly between only 0.5 and 1. Such an effect is dwarfed by the intensive margin elasticity ($\sigma - 1$) with likely values between 4 and 9 (if σ ranges between 5 and 10), under certain assumptions. The key assumption pertains to the effect of an EIA on fixed costs versus variable costs. Suppose $\ln \tau_{ijt} = \rho_0 - \rho_1 \ln \text{DIST}_{ij} - \rho_2 \text{EIA}_{ijt} + \mu_{ijt}$, as is conventional (μ_{ijt} denoting a random error term). Suppose also $\ln f_{ijt} = \delta_0 - \delta_1 \ln \text{DIST}_{ij} - \delta_2 \text{EIA}_{ijt} + \phi_{ijt}$ (ϕ_{ijt} denoting a random error term). If $\delta_2 > \rho_2$ by a large amount, the effect of f_{ijt} changes on the extensive margin may be sufficient to cause the extensive margin trade elasticity to exceed the intensive margin trade elasticity. To date, to our knowledge, there are no firm estimates of δ_2 because there is no data on $\ln f_{ijt}$.

¹¹ Partial (or direct) effects ignore general-equilibrium (or indirect) effects. While techniques exist for estimating the indirect effects, such estimation is beyond this paper's scope, cf., Anderson (2011), Bergstrand and Egger (2011), and Egger et al. (2011) on these issues.

run, as well as offers an alternative approach to instrumental variables using cross-sectional data (and potentially avoids possible shortcomings of the latter approach).¹²

Given the problems associated with accounting for endogeneity of EIAs using instrumental variables and cross-section data, BB argued that a better approach to eliminate endogeneity bias of EIAs is to use panel techniques. In the context of the theory and endogenous self-selection of country pairs into EIAs, BB argued that one method to obtain consistent estimates of the partial effect of EIAs is by fixed effects estimation of:

$$\ln X_{ijt} = \beta_0 + \beta_1 (EIA_{ijt}) + \eta_{ij} + \delta_{it} + \psi_{jt} + \epsilon_{ijt} \quad (2)$$

where η_{ij} is a country-pair fixed effect to capture all time-invariant unobservable bilateral factors influencing nominal trade flows and δ_{it} and ψ_{jt} are exporter-time and importer-time fixed effects, respectively, to capture time-varying exporter and importer GDPs as well as all other time-varying country-specific unobservables in i and j influencing trade, including the exporter's and importers' "multilateral price/resistance" terms (cf., Anderson and van Wincoop, 2003). We refer to this as the fixed-effects (FE) specification. It is important to note that, in most gravity-equation applications using a comprehensive set of RHS variables, the vast bulk of "bilateral" trade-cost variables are time-invariant, such as bilateral distance, common border, common language, etc. BB showed that the partial effect of the typical EIA on nominal trade flows was about 0.76, implying that the typical EIA increased bilateral trade by about 114% after 10–15 years.

BB also employed an alternative specification using first-differencing:

$$\Delta_5 \ln X_{ijt} = \beta_0 + \beta_1 (\Delta_5 EIA_{ijt}) + \delta_{5,it} + \psi_{5,jt} + v_{5,ijt} \quad (3)$$

where Δ_5 refers to first-differencing over 5 years. We refer to this as the first-difference (FD) specification. Note that the bilateral country-pair fixed effects are eliminated; however, the exporter-time ($\delta_{5,it}$) and importer-time ($\psi_{5,jt}$) fixed effects are retained to capture changes in the time-varying exporter and importer GDPs and multilateral price terms over the same 5-year period. The latter effects were ignored in Foster et al. (2011), creating potential omitted variables bias.

First-differencing the panel data yields some potential advantages over fixed effects.¹³ First, it is quite plausible that the unobserved factors influencing the likelihood of an EIA (say, trade below its "natural" level) are likely slow moving and hence serially correlated. If the ϵ_{ijt} are highly serially correlated, the inefficiency of FE is exacerbated as T gets large. This suggests that differencing the data will increase estimation efficiency for our large- T panel. Second, aggregate trade flow data and real GDP data are likely "close to" unit-root processes. Using FE is equivalent to differencing data around the mean (in our sample, year 1980); this may create a problem since T is large in our panel. As Wooldridge (2000, p. 447) notes, if the data follow unit-root processes and T is large, the "spurious regression problem" can arise in a panel using FE.

¹² As argued in BB, the problem with using cross-section data and consequently having to employ IV techniques to account for EIA selection bias is the inability practically of satisfying the "exclusion restriction" with confidence. Most variables that influence trade flows also explain selection into EIAs, and it is difficult to find a variable that explains EIAs that does not also explain trade flows. Egger et al. (2011) used the IV approach to account for the endogeneity of EIAs in their single cross-section, also allowing for selection into zeros trade (i.e., a bivariate probit model). They found using an approach similar to Helpman et al. (2008), except also allowing for endogenous EIAs, that EIAs predominantly affected trade at the (country) intensive margin.

¹³ As Wooldridge (2010, Ch. 10) notes, when the number of time periods (T) exceeds two, the FE estimator is more efficient under the assumption of serially uncorrelated error terms ϵ_{ijt} . The FD estimator is more efficient (when $T > 2$) under the assumption that the error term ϵ_{ijt} follows a random walk (i.e., that the error term $v_{5,ijt} = \epsilon_{ijt} - \epsilon_{ijt-5}$ is white noise). When the number of time periods is exactly two ($T = 2$), estimation with FE and FD produce identical estimates and inferences; then, FD is easier to estimate. When $T > 2$, the choice depends upon the assumption one wants to make about the distribution of the error term ϵ_{ijt} .

FD yields data that deviates from the previous period of our panel, and thus is closer to a unit-root process. In the following, we focus on a first-difference approach; however, the FE estimates are provided in the online appendix to this paper.¹⁴

In this paper, we introduce one further innovation relative to BB.¹⁵ While changes over time in exporter-specific and importer-specific unobservable variables are captured with δ_5 and ψ_5 , changes over time in pair-specific unobservables, such as falling variable and fixed export costs unrelated to EIAs, are not accounted for. Wooldridge (2000) suggests a "random growth" (also called, random trend) first-difference model, henceforth, RGFD model. Unobservable pair-specific changes over time can be partially accounted for by including pair-specific ij fixed effects in Eq. (3), suggesting specification:

$$\Delta_5 \ln X_{ijt} = \beta_0 + \beta_1 (\Delta_5 EIA_{ijt}) + \delta_{5,it} + \psi_{5,jt} + \eta_{ij} + v_{5,ijt}. \quad (4)$$

Consequently, if unobservable declines in bilateral variable and fixed trade costs (say, due to technological improvements) evolve smoothly over time, the η_{ij} s in Eq. (4) will account for these influences. The incorporation of such fixed effects (η_{ij} s) in a first-difference specification was used in Trefler (2004), with such effects referred to there as "secular growth" controls.

One of the other potential contributions of BB's panel methodology was to show that the full impact of EIAs on trade flows took 10–15 years. One reason is that most EIAs are "phased-in" over 5–10 years. The second reason is the lagged effect of the trade-cost changes (such as terms-of-trade changes) on trade flows. As in BB, using a panel allows for differentiating the shorter-term effects (5 years) from the longer-term effects (5–10 years). In the context of the recent developments in the trade literature emphasizing intensive versus extensive margin effects, our panel approach allows for differential timing of these effects. In reality, one would expect that the intensive margin would be affected by a trade-cost change sooner than the extensive margin, because intensive margin changes in volumes do not require any startup costs. Such costs – critical to the extensive margin – may delay the entry of new firms into exporting, and thus we should expect the intensive margin to be influenced in the shorter term and the extensive margin in the longer term, as the results in Bernard et al. (2009) show. Our panel data approach allows for evaluating this hypothesis.¹⁶

It will be useful now to rationalize the use of 5-year differencing of data as in BB, rather than, say, annual differencing. Cheng and Wall (2005) and Wooldridge (2000) both argue in favor of using data differenced over a longer period than annually. Cheng and Wall (2005, p. 8) note 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." Wooldridge (2000, p. 423) confirms the reduction in standard errors of coefficient estimates using changes over longer periods of time than using "year-to-year" changes. Based upon these considerations, we chose as in BB 5-year differences; similar considerations led to the use of 4-year differences in Anderson and Yotov (2011).

¹⁴ It turns out that, for FTAs and deeper EIAs, the results using FE and FD are quite similar. As a practical matter, the choice is more important for TWPTAs and OWPTAs; we discuss this later.

¹⁵ We thank a referee and the editor for motivating this innovation.

¹⁶ These differential timing effects were ignored in Foster et al. (2011). As discussed earlier, two recent theoretical papers suggest a reason for the low trade-cost elasticity of trade flows in macroeconomic analyses using time-series data and the relatively higher trade-cost elasticities of trade in cross-sectional trade analyses. Ruhl (2008) explains this puzzle by noting that the macroeconomic time-series approach is estimating the intensive margin effect of trade, whereas the trade literature's cross-sectional approach is capturing the intensive and extensive margin effects, due to export fixed costs for new producers delaying trade effects and entry. In a complementary approach, Arkolakis et al. (2011) present a demand-oriented staggered-adjustment "Calvo-pricing" approach to explain the lower time-series elasticity in terms of solely an intensive margin effect, and the higher long-run cross-section trade-cost elasticity capturing the longer-term extensive margin elasticity also.

Nevertheless, in a sensitivity analysis later, we will confirm our findings using annual data.

BB did not estimate differential effects of various types of EIAs (in terms of depth of integration) on trade flows. Magee (2008) and Roy (2010) using the methodology of BB found that trade flows were impacted by larger amounts for customs unions relative to FTAs. However, no empirical study has examined the differential impact of FTAs relative to deeper EIAs on the extensive versus intensive margins – much less the differential timing of such effects; these are goals of this paper.¹⁷ The next section discusses how we decompose data into the two margins.

2.3. The Hummels–Klenow margin-decomposition methodology

Hummels and Klenow (2005), or HK, was the first paper to highlight a tractable method for decomposing transparently the extensive and intensive goods margins of trade for a large set of countries' bilateral trade flows using publicly available disaggregate trade data.¹⁸

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

$$EM_{ijt} = \frac{\sum_{m \in M_{ij}} X_{Wjt}^m}{\sum_{m \in M_{Wj}} X_{Wjt}^m} \quad (5)$$

where X_{Wjt}^m is the value of country j 's imports from the world in product m in year t , M_{Wjt} is the set of all products exported by the world to j in year t , and M_{ijt} is the subset of all products exported from i to j in year t . Hence, EM_{ijt} is a measure of the fraction of all products that are exported from i to j in year t , where each product is weighted by the importance of that product in world exports to j in year t .¹⁹

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

$$IM_{ijt} = \frac{\sum_{m \in M_{ijt}} X_{ijt}^m}{\sum_{m \in M_{ij}} X_{Wjt}^m} \quad (6)$$

where X_{ijt}^m is the value of exports from i to j in product m in year t . Thus, IM_{ijt} represents the market share of country i in country j 's imports from the world within the set of products that i exports to j in year t .

¹⁷ It is useful to note here a parallel literature examining the effect of GATT and/or WTO membership on trade flows. For brevity, we note that there now appears little convincing evidence of substantive GATT/WTO effects on trade, once one accounts for EIA dummies, multilateral resistance, and unobserved country-pair fixed effects (as we do here). This is the conclusion of Eicher and Henn (2011) (though they found a non-trivial WTO “terms-of-trade” effect) and of Felbermayr and Kohler (2010) who examined possible extensive margin effects; Eicher and Henn (2011) ignored extensive versus intensive margin effects. We also note an issue raised in Martin and Ng (2004), which is the role of multilateral tariff reductions under the GATT/WTO. Most-Favored-Nation (MFN) tariff cuts could also be affecting results. However, such MFN tariff cuts by country would be accounted for by the exporter-time and importer-time fixed effects.

¹⁸ Studies have also used country-specific data on individual plants (or firms) to study the extensive and intensive firm margins of trade liberalization, but such studies have necessarily been confined to particular countries because such data is widely known to be much more costly to access and such data sets have not been concorded for international comparisons, as noted in Helpman et al. (2008). See Eaton et al. (2008) for a study of French firms, Trefler (2004) for a study of Canada and the United States, and Pavcnik (2002) for a study of Indian firms. Another relevant theoretical and empirical piece with similar overtones is Arkolakis et al. (2008).

¹⁹ Alternatively, one could use an unweighted average, which would then be simply the fraction of all products exported from i to j . However, HK – as well as researchers since then – use the weighted average. A weighted average seems more appropriate since cars and pencils do not have the same values in trade. Also, since we will use a time series of cross sections, we will consider later two alternative methods for fixing the trade-share weights over time.

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

$$EM_{ijt} IM_{ijt} = \frac{\sum_{m \in M_{ij}} X_{ijmt}}{\sum_{m \in M_{Wj}} X_{Wjmt}} = X_{ijt} / X_{jt} \quad (7)$$

where X_{jt} denotes j 's imports from the world. Taking the natural logs of Eq. (7) and some algebra yields:

$$\ln X_{ijt} = \ln EM_{ijt} + \ln IM_{ijt} + \ln X_{jt}. \quad (8)$$

Consequently, the HK decomposition methodology yields that the log of the value of the trade flow from i to j in any year t can be decomposed linearly into (logs of) an extensive margin, an intensive margin, and the value of j 's imports from the world. We note three issues regarding the HK methodology. First, since we will focus empirically on the RGFD specification similar to Eq. (4), the term $\Delta_5 \ln X_{jt}$ will be subsumed in the importer-time fixed effect $\psi_{5,jt}$. Second, HK applied their methodology to only a cross section. By contrast, we are applying it to a time series of cross sections. Consequently, the trade weights used in constructing EM_{ijt} and IM_{ijt} will likely vary from year to year. To address this, we will also consider later in a sensitivity analysis fixed-year trade-share weights and also a chain-weighting technique. Third, there are numerous zeros in the variables in Eq. (8) and the results may be biased by ignoring the existence of firm heterogeneity. Hence, we will address later why our panel approach largely alleviates sample-selection bias and firm-heterogeneity bias, as raised in Helpman et al. (2008).²⁰

Finally, we are distinguishing between various types of EIAs and allow for lagged effects. Hence, the actual specification for the RGFD versions of our model with no lags is:

$$\Delta_5 \ln X_{ijt} = \beta_0 + \beta_1 (\Delta_5 CUCMECU_{ijt}) + \beta_2 (\Delta_5 FTA_{ijt}) + \beta_3 (\Delta_5 TWPTA_{ijt}) + \beta_4 (\Delta_5 OWPTA_{ijt}) + \delta_{5,it} + \psi_{5,jt} + \eta_{ij} + v_{5,ijt} \quad (9)$$

$$\Delta_5 \ln EM_{ijt} = \theta_0 + \theta_1 (\Delta_5 CUCMECU_{ijt}) + \theta_2 (\Delta_5 FTA_{ijt}) + \theta_3 (\Delta_5 TWPTA_{ijt}) + \theta_4 (\Delta_5 OWPTA_{ijt}) + \delta_{5,it} + \psi_{5,jt} + \eta_{ij} + v_{5,ijt} \quad (10)$$

$$\Delta_5 \ln IM_{ijt} = \lambda_0 + \lambda_1 (\Delta_5 CUCMECU_{ijt}) + \lambda_2 (\Delta_5 FTA_{ijt}) + \lambda_3 (\Delta_5 TWPTA_{ijt}) + \lambda_4 (\Delta_5 OWPTA_{ijt}) + \delta_{5,it} + \psi_{5,jt} + \eta_{ij} + v_{5,ijt}. \quad (11)$$

In specifications including lagged EIA variables, we will use, for instance, the notation $Lag_{\Delta_5} EIA_{ijt}$ to denote an EIA formed 5 to 10 years prior to the trade-flow change and $Lead_{\Delta_5} EIA_{ijt}$ to denote an EIA formed in the 5 years prior to the trade-flow change. Recall that $OWPTA$ denotes a one-way preferential trade agreement, $TWPTA$ denotes a two-way preferential trade agreement, FTA denotes a free trade agreement, and $CUCMECU$ denotes “deeper” EIAs, defined just below.

Consequently, the literature to date suggests that the endogeneity of EIAs in typical gravity equations may bias the estimation of partial effects of EIAs on trade flows and trade margins. We augment the panel approach in BB to account for random growth in bilateral trade flows and the margins of trade due to unobservable changes in bilateral

²⁰ Besedes and Prusa (2011) emphasize that extensive margin (intensive margin) effects may be overstated (understated) in examining effects of liberalizations using panel data, due to “survival” issues. Addressing this issue is beyond the scope of this particular paper. However, our results may not be biased excessively by this issue since we find material intensive margin effects from EIAs, unlike Foster et al. (2011).

variable and fixed export costs that evolve smoothly over time. We follow the methodology established in HK to decompose trade flows (log-) linearly into extensive and intensive margins. In the next section, we estimate equations such as Eqs. (9)–(11) above to determine if both intensive and extensive (goods) margins are affected by EIAs, the differential effects of various “types” of EIAs on these extensive and intensive margins of trade, and the differential “timing” of the partial effects of EIAs on the two margins.

3. Data

The two key variables for our empirical analysis are disaggregate bilateral trade flows and a multichotomous index of the level of economic integration agreement (EIA) between a large number of country pairs for a large number of years.

First, while several earlier gravity-equation analyses have used dummy variables indicating the presence or absence of an EIA between country pairs for numerous years, such as Rose (2004), there are few publicly available systematic data sets that have multichotomous indexes of EIAs for a large number of country pairs and number of years (a panel). We use the data set constructed by Scott Baier and Jeffrey Bergstrand and provided at Jeffrey Bergstrand's website, www.nd.edu/jbergstr/.²¹ The index is defined as: no EIA (0), one-way preferential trade agreement, or *OWPTA* (1), two-way preferential trade agreement *TWPTA* (2), free trade agreement, or *FTA* (3), customs union (4), common market (5), and economic union (6). The definitions are conventional, based upon Frankel (1997), and are defined explicitly in the data set. Because of the small number of “deeper EIAs,” we combined customs unions (4), common markets (5), and economic unions (6) into one variable, *CUCMECU*. One of the strengths of the Baier–Bergstrand EIA panel is, for 98.6% of the cells where the EIA status of the country pair changes (from 0 to 1, 0 to 3, 2 to 3, etc.), there exists a hyperlink to a copy (PDF format) of the original treaty. Because of the extensive number of bilateral and plurilateral EIAs of differing degrees of depth, the difficulty of listing all such agreements precludes a table; we recommend downloading the zip file from the web site.

Second, annual bilateral trade flows for 1962–2000 are from the NBER–United Nations trade data set at www.nber.org/data and documented in Feenstra et al. (2005). The data are organized by 4-digit SITC, Revision 2. It covers trade flows reported by 149 countries and covers 98% of world exports. This is the most disaggregated publicly available data set for bilateral trade flows for a large number of years and a large number of country pairs, constructed on a consistent basis, necessary for the analysis at hand. This 4-digit SITC data was also used in Hillberry and McDaniel (2002), Kehoe and Ruhl (2009), and Foster et al. (2011). In 1962, this resulted in 969 categories of “goods”; in 2000, this resulted in 1,289 categories of goods. One concern is that the level of disaggregation is not high enough, biasing results toward the intensive margin, as discussed transparently in HK; as one aggregates up, there become fewer categories in which a country does not export. Other studies have used higher levels of disaggregation to calculate trade flows in a good to avoid this (intensive margin) aggregation bias. For instance, Broda and Weinstein (2006) used 7-digit data for the years 1972–1988 and then 10-digit data for the years 1990–2001. However, this was U.S. import data only, and for two short panels. Constraining ourselves to U.S. import trade flows only is problematic because the United States has only a small number of FTAs and has no deep EIAs, precluding evaluating our hypotheses. Since our decomposition of the extensive and intensive margins is based upon HK, we compare our data set to theirs. First, HK used only a cross section of UNCTAD TRAINS data for the year 1995, but for a large number of country pairs like

here.²² Using a cross section is problematic because we are relying upon panel techniques to avoid endogeneity bias. Second, HK used data at the Harmonized System 6-digit classification code; this yielded 5017 goods categories, five times our number of categories. However, HK examined the levels of extensive margins and the correlations between the extensive margin and factors influencing the extensive margin in cross-sectional data – GDP, employment (L), GDP/L – at various levels of aggregation (1-digit, 2-digit, ..., 6-digit) and found two interesting results. As expected, extensive margins were much lower, and the correlation between the extensive margin of trade and determinants of it (such as GDP) were lower, as data became more aggregated. For example, at the 6-digit level, 62% of the GDP elasticity of exports could be explained by the extensive margin, whereas at the 1-digit level only 11% of the elasticity could be explained by the extensive margin. However, interestingly, at the 4-digit level, a sizable 54% of this elasticity could still be explained by the extensive margin. For the per capita GDP elasticity of exports, 66 (62) percent of this elasticity was explained by the extensive margin at the 6-digit (4-digit) level. These results suggest that the (intensive margin) aggregation bias may not be that severe in using 4-digit categories, rather than the 6-digit categories in HK.

In a previous section, we provided an econometric justification for using 5-year differencing of the data (beginning in 1970 and ending in 2000, i.e., the first 5-year period is 1965–1970). Nevertheless, in a sensitivity analysis later, we will confirm our empirical findings using annual data. Due to space constraints of the journal, we focus in this paper on the random growth first-difference (RGFD) results. However, FD results (without random-growth trends) and fixed effects results in levels are reported in an online appendix.

4. Empirical results

The main empirical results for Eqs. (9)–(11) are presented in Table 1 in Section 4.1. In Section 4.2, Table 2 reports another set of RGFD results using an alternative chain-weighted technique for weights, due to the estimation of a time series of cross sections. In Section 4.3, we discuss the role of potential endogeneity bias due to country selection and firm heterogeneity. We address how our panel techniques largely alleviate these biases, raised in Helpman et al. (2008). In Section 4.4, we address the sensitivity of the findings to using instead annual data and show why the results are largely insensitive to reverse causality bias.

Within each set are the results of running the same specification for three alternative LHS variables. TRADE refers to the aggregate bilateral trade flow from i to j (or X_{ijt} in Eq. (8)). EM refers to the extensive margin (or EM_{ijt} in Eq. (8)). IM refers to the intensive margin (or IM_{ijt} in Eq. (8)). From Eq. (8), once changes in $\ln X_{ijt}$ are controlled for with importer-year fixed effect ψ , the sum of variations in the extensive and intensive margins must equal the variation in the aggregate trade flow. This allows us to use the empirical results to infer the relative extensive and intensive margin elasticities to a trade liberalization.

4.1. Main results

Table 1 reports the results using RGFD Eqs. (9)–(11) with no lag in Set 1 and with one lag in Set 2.²³ Our first set of results in Set 1 is promising; we highlight the key issues. First, we note in column (1a) that deeper EIAs have larger effects on aggregate trade flows than FTAs, and the latter have larger effects on trade than two-way or one-way PTAs, as expected. The coefficient estimates for $\Delta_5 CUCMECU_{ijt}$ and for $\Delta_5 FTA_{ijt}$ are economically and statistically significant. Second, for deep EIAs, FTAs, and one-way PTAs, the intensive-margin effects in

²¹ This data set was constructed under National Science Foundation grants SES-0351018 and SES-0351154 and includes annually from 1960 to 2005 for the pairings of 195 countries an index ranging from 0 to 6 of the level of any EIA between the pair.

²² For robustness, HK also examined a cross section of U.S. imports only at a higher level of disaggregation. But this raises the same problem as in Broda and Weinstein (2006) of U.S. data only.

²³ FD and FE results are in the online appendix.

Table 1
5-Year differenced data.

| Variables | Set 1 (RGFD) | | | Set 2 (RGFD) | | |
|---------------------------------------|----------------------------|-------------------------|-------------------------|----------------------------|-------------------------|-------------------------|
| | (1a) | (1b) | (1c) | (2a) | (2b) | (2c) |
| | $\Delta_5 \ln TRADE_{ijt}$ | $\Delta_5 \ln EM_{ijt}$ | $\Delta_5 \ln IM_{ijt}$ | $\Delta_5 \ln TRADE_{ijt}$ | $\Delta_5 \ln EM_{ijt}$ | $\Delta_5 \ln IM_{ijt}$ |
| $\Delta_5 CUCMECU_{ijt}$ | 0.329*** (0.072) | 0.100* (0.062) | 0.229*** (0.065) | 0.387*** (0.077) | 0.106* (0.067) | 0.281*** (0.069) |
| Lag $\Delta_5 CUCMECU_{ijt}$ | | | | 0.309*** (0.076) | 0.131** (0.065) | 0.179*** (0.072) |
| $\Delta_5 FTA_{ijt}$ | 0.192*** (0.053) | 0.074* (0.043) | 0.118*** (0.046) | 0.242*** (0.056) | 0.085** (0.10046) | 0.157*** (0.049) |
| Lag ΔFTA_{ijt} | | | | 0.228*** (0.055) | 0.084** (0.047) | 0.145*** (0.049) |
| $\Delta_5 TWPTA_{ijt}$ | −0.001 (0.072) | 0.012 (0.055) | −0.013 (0.063) | 0.069 (0.079) | 0.023 (0.063) | 0.046 (0.068) |
| Lag $\Delta_5 TWPTA_{ijt}$ | | | | 0.111* (0.071) | 0.071 (0.063) | 0.041 (0.068) |
| $\Delta_5 OWPTA_{ijt}$ | 0.072 (0.065) | −0.064 (0.053) | 0.135*** (0.062) | 0.116** (0.067) | −0.043 (0.055) | 0.159*** (0.063) |
| Lag $\Delta_5 OWPTA_{ijt}$ | | | | 0.285*** (0.069) | 0.171*** (0.055) | 0.114** (0.067) |
| Constant | 0.413*** (0.120) | 0.167* (0.130) | 0.247* (0.182) | 0.608*** (0.247) | −0.089 (0.274) | 0.697** (0.342) |
| Fixed effects | | | | | | |
| Exporter-year ($i, t - (t - 5)$) | Yes | Yes | Yes | Yes | Yes | Yes |
| Importer-year ($j, t - (t - 5)$) | Yes | Yes | Yes | Yes | Yes | Yes |
| Country-pair (ij) | Yes | Yes | Yes | Yes | Yes | Yes |
| R ² | 0.294 | 0.365 | 0.357 | 0.330 | 0.402 | 0.403 |
| No. of observations | 48,619 | 48,619 | 48,619 | 41,767 | 41,767 | 41,767 |

Notes: Robust standard errors in parentheses. *, **, and *** denote statistical significance at the 10, 5, and 1% levels, respectively, in two-tailed t-tests.

column (1c) are economically and statistically significant. The effect is largest for $\Delta_5 CUCMECU_{ijt}$, next largest for $\Delta_5 OWPTA_{ijt}$, and slightly smaller for $\Delta_5 FTA_{ijt}$; the $\Delta_5 TWPTA_{ijt}$ effect is trivially small. The extensive margin effects in column (1b) are statistically significant for deep EIAs and FTAs; moreover, the extensive margin effect for $\Delta_5 CUCMECU_{ijt}$ is larger than that for $\Delta_5 FTA_{ijt}$, as expected. Third, we note that – in the absence of lagged effects – for deep EIAs, FTAs, and OWPTAs the intensive margin effect is *always larger* than its corresponding extensive margin effect, consistent with our interpretation of the Melitz model in Chaney (2008). Thus, the Set 1 results using specifications (9)–(11) provide substantive support for three major hypotheses. However, by adding a lagged change in the EIAs, we can even provide evidence for a fourth hypothesis, extensive (intensive) margin effects should become relatively more (less) important with time.

Even more interesting are the results in Set 2 where we allow current and lagged changes, consequently capturing the effects on trade and the margins of 10 years of EIA changes. As discussed earlier and in BB, EIAs are likely to have delayed impacts on trade flows. One reason is that EIAs are typically “phased-in” over 5 to 10 years, delaying the full implementation of liberalization. Second, the impact on trade flows of EIAs works largely through terms-of-trade effects. It is well known that terms-of-trade changes can also have a delayed impact on trade flows. BB found that most of the impact of EIAs on aggregate trade flows was captured using panel techniques in a period of 10 years. Consider deep EIAs in columns (2a)–(2c). First, for aggregate trade, consistent with earlier results, common membership in a deep EIA increases two members' bilateral trade flow by a plausible 101% after 10 years ($e^{0.387 + 0.309} = 2.01$). Second, we find that in the short term, the effect on trade comes primarily through the intensive margin, as expected. The intensive-margin effect (0.281) dominates the extensive-margin effect (0.106) for $\Delta_5 CUCMECU_{ijt}$, with both effects statistically significant. However, for the lagged change the extensive margin effect (0.131) is statistically significant and larger than the current-period extensive margin effect. Also, while the lagged intensive margin effect (0.179) is statistically significant, it is smaller than the

current-period intensive margin effect (0.281). Third, when we sum the current period and lagged effects for each margin, we find an extensive margin effect of 0.237 and an intensive margin effect of 0.460, consistent with our theoretical conjecture on the relative size of intensive margin and extensive margin elasticities.

Consider now FTAs. First, for aggregate trade, common membership in an FTA increases two members' bilateral trade flow by 60% after 10 years ($e^{0.242 + 0.228} = 1.60$). The intensive margin effect dominates the extensive margin effect in both current period and lagged period. The extensive margin effect does not increase over time; however, in a robustness analysis later for the FD specification, we find the relative size of the extensive margin effect increasing in the lag. Second, when we sum the current-period and lagged effects for each margin, we find an extensive margin effect of 0.169 and an intensive margin effect of 0.302. Hence, the ratio of these effects is similar to that for *CUCMECU*, with the intensive margin elasticity larger than the extensive margin elasticity as expected.

For two-way PTAs, we found in Set 2 only a statistically significant effect of the lagged $\Delta_5 TWPTA$ change on $\Delta_5 \ln TRADE_{ijt}$. We found only small positive but not statistically significant effects on the two margins. Nevertheless, the relative effects for $\Delta_5 TWPTA_{ijt}$ conformed to earlier results. For the short term, the relatively larger effect was at the intensive margin (0.046 for intensive relative to 0.023 for extensive). For the lagged change, the relatively larger effect was at the extensive margin (0.071 for extensive relative to 0.041 for intensive).

For one-way PTAs, we found in the current period the intensive margin effect (0.159) was economically and statistically significant and dominated the statistically insignificant extensive-margin effect (−0.043). For the lagged effect, the extensive margin effect dominated the intensive margin effect as expected. In fact, the extensive margin effect of 0.171 was statistically significant and the intensive margin effect of 0.114 was also statistically significant.

Our evidence indicates that the four hypotheses posed earlier generally hold. Both extensive and intensive margins are affected by EIAs, deeper EIAs tend to have larger impacts on trade, the extensive margin,

and the intensive margin than shallower agreements, and intensive margin effects tend to occur sooner than extensive margin effects. Moreover, total intensive margin elasticities tend to be larger than total extensive margin elasticities, consistent with our fourth hypothesis suggested by the Melitz model in Chaney (2008) and empirical estimates of $\gamma/(\sigma - 1)$. In the next sections, we evaluate the robustness of these results.

4.2. Sensitivity analysis 1: chain-weighting

The original HK analysis of extensive and intensive margins was conducted using a cross section for a particular year (using two alternative data sets). Our analysis uses a time series of cross sections. Consequently, the variables EM_{ijt} and IM_{ijt} defined in Eqs. (5) and (6), respectively, include trade “weights” that change over time. Consequently, we consider two alternative methods for holding constant the trade weights, X_{Wjmt} , over time.

The first alternative measure of the extensive margin, denoted EM_{ijt}^* , uses values for X_{Wjmt} set to a “base year”; we chose 1995. Hence, EM_{ijt}^* is defined as:

$$EM_{ijt}^* = \frac{\sum_{m \in M_{ijt}} X_{Wjmt1995}}{\sum_{m \in M_{ijt}} X_{Wjmt1995}} \quad (12)$$

where $X_{Wjmt1995}$ is the value of country j 's imports from the world in product m in year 1995. M_{Wjt} and M_{ijt} are defined as before. In order to ensure that the log-linear decomposition holds analogous to Eqs. (7) and (8), we define the intensive margin of goods exported from i to j as:

$$IM_{ijt}^* = \frac{\sum_{m \in M_{ijt}} X_{ijmt}}{\sum_{m \in M_{ijt}} X_{Wjmt1995}} \quad (13)$$

Consequently, the product of EM_{ijt}^* and IM_{ijt}^* results in:

$$\ln X_{ijt} = \ln EM_{ijt}^* + \ln IM_{ijt}^* + \ln X_{j1995} \quad (14)$$

In estimation (using log-levels), the last term on the RHS is subsumed in a fixed effect.

One problem with the fixed-year trade weights is that the particular year chosen may bias the results, especially given that our data set spans 1965–2000. To address this, we considered another trade-weight approach. The second alternative measure for our RGFD specification uses a “chain-weighted” approach. Since the LHS variables are 5-year differences, we used the same chain-weighting technique as used in National Income Accounts. For the X_{Wjmt} , we used for t the geometric average of the trade flows for the corresponding years.

Table 2 presents in Set 3 the results of using the chain-weighted trade weights; consequently, the results for Set 3 can be compared readily to the previous identical Set 2 specifications (which use the time-varying weights).²⁴ For brevity, we report only the result for the RGFD specification including one lag. Of course, the alternative weighting approach has no bearing on the aggregate trade flow regressions. It is readily seen that there is no material difference between Set 2 and Set 3 results for the extensive and intensive margins.

4.3. Sensitivity analysis 2: accounting for country-selection and firm-heterogeneity biases

First, as is well known, most bilateral trade flows among a large number of country pairs include numerous observations with “zeros.” As noted in both Helpman et al. (2008), or HMR, and Egger et al.

Table 2
5-Year differenced data (chained).

| Variables | Set 3 (RGFD-Chained) | | |
|---------------------------------------|----------------------------|-------------------------|-------------------------|
| | (3a) | (3b) | (3c) |
| | $\Delta_5 \ln TRADE_{ijt}$ | $\Delta_5 \ln EM_{ijt}$ | $\Delta_5 \ln IM_{ijt}$ |
| $\Delta_5 CUCMECU_{ijt}$ | 0.387*** (0.077) | 0.113** (0.068) | 0.274*** (0.070) |
| Lag $\Delta_5 CUCMECU_{ijt}$ | 0.309*** (0.076) | 0.124** (0.067) | 0.185*** (0.073) |
| $\Delta_5 FTA_{ijt}$ | 0.242*** (0.056) | 0.090** (0.047) | 0.152*** (0.050) |
| Lag $\Delta_5 FTA_{ijt}$ | 0.228*** (0.055) | 0.076* (0.048) | 0.153*** (0.050) |
| $\Delta_5 TWPTA_{ijt}$ | 0.069 (0.079) | 0.025 (0.064) | 0.044 (0.069) |
| Lag $\Delta_5 TWPTA_{ijt}$ | 0.111* (0.071) | 0.053 (0.064) | 0.058 (0.069) |
| $\Delta_5 OWPTA_{ijt}$ | 0.116** (0.067) | −0.051 (0.055) | 0.168*** (0.064) |
| Lag $\Delta_5 OWPTA_{ijt}$ | 0.285*** (0.069) | 0.168*** (0.069) | 0.117** (0.068) |
| Constant | 0.608*** (0.247) | −0.066 (0.278) | 0.679** (0.348) |
| Fixed effects | | | |
| Exporter-year ($i, t - (t - 5)$) | Yes | Yes | Yes |
| Importer-year ($j, t - (t - 5)$) | Yes | Yes | Yes |
| Country-pair (ij) | Yes | Yes | Yes |
| R ² | 0.330 | 0.407 | 0.407 |
| No. of observations | 41,767 | 41,767 | 41,767 |

Notes: Robust standard errors in parentheses. *, **, and *** denote statistical significance at the 10, 5, and 1% levels, respectively, in two-tailed t-tests.

(2011), such zeros can be associated with selection bias (often referred to as “selection into exporting”); this can arise from the existence of fixed exporting costs and may be associated with firm heterogeneity in productivities (but does not require it). Second, even in the absence of zero trade flows, the existence of firm heterogeneity may bias our results. Hence, our results may be sensitive to absence of controls for sample-selection and firm-heterogeneity biases. While one option is to adapt the cross-sectional approach of HMR to our panel setting (which we actually do and present in the online appendix), it is unnecessary due to the random-growth first-difference (RGFD) approach we use. We explain in the context of a representative gravity equation generated from a Melitz model, such as Eq. (1). In this case, there are fixed export costs and firm heterogeneity. The key issue is the existence of these two factors allows selection of firms in country i as exporters into destination market j . In our context, let Z_{ijt}^m be a latent variable reflecting the ratio of variable export profits to fixed export costs for the most productive firm in country i in year t (a_{ijt}^m); positive exports from i to j occur if $Z_{ijt}^m > 1$. As discussed in HMR, coefficient estimates in an (aggregate) trade flow gravity equation need to control for variation in Z_{ijt}^m ; HMR show that accounting for Z_{ijt}^m controls both for Heckman selection bias (because the inverse Mills' ratio is a monotonic function of Z_{ijt}^m) and for firm-heterogeneity bias (with a control that is a function of both Z_{ijt}^m and the inverse Mills' ratio (which is a function of Z_{ijt}^m)). Hence, for our purposes, we need to account for fluctuations in Z_{ijt}^m across country pairs and over time. However, this is the purpose of using the random-growth first-difference model. If the factors influencing selection and firm heterogeneity evolve smoothly over time, the RGFD model will account for the controls used in HMR. Unlike the cross-sectional context of HMR and Egger et al. (2011), we use first-differences to eliminate any unobservable differences between country pairs in the time-invariant components of τ_{ijt}^m , f_{ijt}^m , and Z_{ijt} , and we use the RGFD model to capture any pair-specific time-varying trends in τ_{ijt}^m , f_{ijt}^m , and Z_{ijt} . Thus, the RGFD model accounts for (unobserved) changes

²⁴ The 1995 base-period approach yielded results not materially different.

across pairs and over time in Z_{ijt}^m .²⁵ Nevertheless, for robustness, we also consider an alternative two-stage approach to control for selection bias and firm heterogeneity in the spirit of HMR to capture changes in the controls that might not be fully accounted for using the RGFD model. Due to space constraints, we can only summarize the results of this sensitivity analysis; the actual results are presented in the online appendix. Importantly, as anticipated based upon the discussion above, the main finding is that there is no material difference in the results after correcting for sample-selection bias and firm heterogeneity using a panel adaptation of the HMR cross-sectional approach. To emphasize, this does not imply that selection bias and firm heterogeneity are absent in the data; our results imply that such biases are largely eliminated due to the first-differencing and pair fixed effects in our RGFD regressions.²⁶

4.4. Sensitivity analysis 3: annual data and reverse causality

Our preferred methodology has been to use 5-year differences of annual data to estimate the partial effects of various types of EIAs on bilateral trade flows, extensive margins, and intensive margins, although estimates using fixed effects are in the online appendix. Based upon Wooldridge (2000, p. 423) and Cheng and Wall (2005, p. 8), we have argued that 5-year differences are more appropriate than annual differences, due to the likelihood that trade flows cannot adjust *within one year* to EIA formations and that time is needed to capture full effects. This is supported by the result that it can take 10–15 years for an EIA to have its full effect. Nevertheless, we have annual data on EIAs, trade flows, and extensive and intensive margins. For robustness, we examine annual data also because we may be able to find even more evidence of the time path of adjustment of trade flows, extensive margins, and intensive margins to EIAs.

In particular, examination of annual data may be useful to better identify the potential role of “reverse causality.” Do EIAs “cause” trade, or does trade “cause” EIAs? This question is relevant because Baier and Bergstrand (2004) showed that the factors that tend to explain bilateral aggregate trade flows also tend to explain the selection of country-pair governments into EIAs. Thus, one might well expect that trade “leads” EIA formations because of this selection effect. However, the key issue here is to try to make sure that these selection effects do not bias our estimates of the partial contemporaneous and lagged effects of EIAs on trade. That is, any leading of trade before EIAs needs to be well in advance of our estimates of EIAs’ effects on trade, needs to be economically small, and needs to diminish as the date of entry into the agreement approaches.

Before proceeding to the annual data, we conducted a simple exogeneity test using the 5-year differenced data as used in Baier and Bergstrand (2007), as suggested in Wooldridge (2010, p. 325). We re-estimated the RGFD specifications in Set 2 to also include either the level of the EIA variables five years forward (e.g., $CUCMECU_{ij,t+5}$) or the change in the EIA variables 5 years before the trade-flow or margin changes (e.g., $CUCMECU_{ij,t+5} - CUCMECU_{ijt}$). We found evidence of positive, statistically significant coefficient estimates of 5-year lead levels of $CUCMECU$ on aggregate trade flows, extensive margins, and intensive margins, of FTA on aggregate trade flows and extensive margins, and of $TWPTA$ on aggregate trade flows and intensive margins.²⁷ Unfortunately, adding further “leads” would further reduce the size of our sample, providing motivation for examining annual data.

As with the 5-year data, we considered both RGFD, FD and FE specifications using annual data. We first examined the relationships using

annual first-differenced data and both the FD and RGFD models. One of the inherent problems with differencing annual data is that the resulting trade-flow, extensive margin, and intensive margin changes are “noisy.” The main result from re-estimating Eqs. (3) and (4) with annual data is that – in general – neither current, lagged, nor lead formations of EIAs had any *systematic* statistically significant effects on annual log-differences in aggregate trade flows, extensive margins, and intensive margins (using up to 7 annual lead and lag changes). The results are presented in the online appendix.

To avoid multicollinearity and the consequent reduced sample size of adding more lead and lag changes, we turned to another approach.²⁸ We introduced linear trends of the lead and lag changes in the EIA variables alongside the concurrent EIA variable. This methodology proved useful, yielding two important insights. Consider the RGFD specification. First, while there were few statistically significant coefficient estimates, the key findings from Set 2 and Set 3 held up. For concurrent $\Delta CUCMECU_{ijt}$ and ΔFTA_{ijt} , their coefficient estimates for the intensive margin were positive (0.115 and 0.075, respectively) and statistically significant at 1%; however, the extensive margin concurrent effects were insignificant. For the lagged trends of $\Delta CUCMECU$ and ΔFTA , their coefficient estimates for the extensive margin were positive (0.034 and 0.033, respectively) and statistically significant; however, the intensive margin effects were insignificant for the lagged trends. The two trends’ annual estimates implied total lagged extensive margin effects of 65% and 63% for $CUCMECU$ and FTA , respectively. All these results made sense. However, the second important finding was that the *lead* trend for $\Delta CUCMECU$ on the extensive margin was an unexpectedly *positive* and statistically significant 0.047.²⁹

Yet, two findings of economically and statistically significant lead effects for $CUCMECU$ on the extensive margin using (5-year and 1-year) first-differenced data suggested further analysis of the annual data using log-levels was warranted. Applying FE specification (2) to annual data with analogous 15-year linear lead and lag annual trends yielded informative results, provided in Table 3.³⁰ We note that Table 3 reports results only for $CUCMECU$ and FTA . The reason is that the FE specifications – whether using all annual data from 1962 to 2000 or only 8 cross-sections five years apart – yielded several negative coefficients for $TWPTA$ and $OWPTA$, whereas the coefficient estimates for $CUCMECU$ and FTA yielded qualitatively and quantitatively similar coefficient estimates using the FE, FD, and RGFD specifications.³¹

Table 3 provides coefficient estimates for $CUCMECU$ and FTA for the linear trend model. The first column of Table 3 lists the two EIAs of interest. $CUCMECU$ refers to coefficient estimates for the concurrent year for

²⁸ Even with only 7 leads and lags, sample size declined from 313,189 to 102,059 observations.

²⁹ The only other statistically significant effects were a positive lag trend effect for aggregate trade flows for ΔFTA and statistically significant negative lead trends for trade flows and extensive margin for $\Delta OWPTA$, which make sense. The full set of results for the RGFD and FD specifications are reported in the online appendix.

³⁰ Note the large number of observations using annual data. Because of the large number of fixed effects in this estimation, we employed the “high dimensional” fixed effects (HD-FE) approach of Abowd and Kramarz (1999). In particular, we extended the two-way high HD-FE iterative approach of Carneiro et al. (2012) to our three-way FE context, clustering the standard errors on the bilateral fixed effects. The procedure employs the notion that the FEs are simple averages. The iterative procedure runs a partitioned regression on the EIAs, the means for each FE are computed, and the regression is re-run using the means as controls, updating at each iteration the mean values until convergence is achieved (with a convergence criteria that the residual sum of squares between the $(n-1)^{th}$ and n^{th} regressions is less than $2.0e-16$.) The approach yields exact least squares solutions for coefficients and standard errors.

³¹ The reason is that the growth of intra-industry trade over this period dominated inter-industry trade growth, and this trend over time caused the coefficient estimates for $TWPTA$ and $OWPTA$ to be *biased downward*; $TWPTAs$ and especially $OWPTAs$ are typically between developed and developing countries. In fact, it was specifically for this reason we focused on the RGFD specification, because the pair fixed effects in the 5-year differenced data accounted for this bias, resulting in positive and some statistically significant coefficients in Sets 1, 2, and 3 in Tables 1 and 2 for $TWPTA$ and $OWPTA$. Nevertheless, the full set of FE coefficient estimates using annual data for 1962–2000 are available in the online appendix.

²⁵ This issue was ignored in Foster et al. (2011) and Eicher and Henn (2011). Also, variation in α_{ij}^m is accounted for by the time-varying exporter fixed effect.

²⁶ In the online appendix, we address that, even though the Religion variable (the first-stage variable providing identification in the second stage) has little variation over time, separate probit cross-section regression coefficient estimates allow identification in the second stage. A possible preferred approach would be a dynamic selection equation, but this approach (as one referee noted) is beyond the scope of the present paper.

²⁷ Using forward changes in EIAs yielded similar results.

Table 3
Annual data in log-levels, 15-year linear trends.

| Variables | Set 4 | | |
|--------------------------------------|--------------------------|-----------------------|-----------------------|
| | (4a) | (4b) | (4c) |
| | $\ln \text{TRADE}_{ijt}$ | $\ln \text{EM}_{ijt}$ | $\ln \text{IM}_{ijt}$ |
| CUCMECU_{ijt} | 0.450*** (0.029) | 0.057** (0.025) | 0.393*** (0.026) |
| $\text{CUCMECU}_{\text{Lag Trend}}$ | 0.025*** (0.003) | 0.010*** (0.003) | 0.015*** (0.003) |
| $\text{CUCMECU}_{\text{Lead Trend}}$ | -0.006*** (0.001) | -0.001 (0.001) | -0.005*** (0.001) |
| FTA_{ijt} | 0.262*** (0.019) | 0.080*** (0.016) | 0.182*** (0.018) |
| $\text{FTA}_{\text{Lag Trend}}$ | 0.028*** (0.002) | 0.025*** (0.002) | 0.004** (0.002) |
| $\text{FTA}_{\text{Lead Trend}}$ | -0.020*** (0.001) | -0.007*** (0.001) | -0.013*** (0.001) |
| Constant | -3.829*** (0.002) | -3.126*** (0.002) | -3.830*** (0.002) |
| Fixed effects | | | |
| Exporter-year (<i>i,t</i>) | Yes | Yes | Yes |
| Importer-year (<i>j,t</i>) | Yes | Yes | Yes |
| Country-pair (<i>ij</i>) | Yes | Yes | Yes |
| R ² | 0.838 | 0.780 | 0.686 |
| No. of observations | 313,189 | 313,189 | 313,189 |

Notes: Robust standard errors in parentheses. *, **, and *** denote statistical significance at the 10, 5, and 1% levels, respectively, in two-tailed t-tests.

CUCMECU, $\text{CUCMECU}_{\text{LagTrend}}$ refers to coefficient estimates on the linear trend for the 15 years of lagged values, and $\text{CUCMECU}_{\text{LeadTrend}}$ refers to coefficient estimates on the linear trend for the 15 years of lead values. Analogous interpretations hold for FTA. The second, third, and fourth columns of numbers report the coefficient estimates for aggregate trade flows, extensive margin, and intensive margin, respectively.

Consider first CUCMECU. The partial effect for CUCMECU (0.450) for aggregate trade flows implies that a CUCMECU increases trade in the current quarter by 57%. The lagged trend for CUCMECU increases trade 2.5% annually, or 45% over 15 years; the combined effect is 102%, which is quite close to the 101% estimated using the 5-year differenced

RGFD results in Table 2. Yet the 15-year lead trend for CUCMECU decreases trade by only 0.6% annually, or 8.6% over 15 years.

For the extensive margin, the partial effect of CUCMECU (0.057) implies that a CUCMECU increases the extensive margin in the current quarter a trivial 6%. However, the lagged trend for CUCMECU is 1% annually, or 16% over 15 years; the combined effect is 22%, which is quite close to the 27% estimated using the 5-year differenced RGFD results in Table 2. The 15-year lead trend for CUCMECU decreases the extensive margin by only 0.1% annually, or 1.5% over 15 years.

For the intensive margin, the partial effect of CUCMECU (0.393) implies a CUCMECU increases the intensive margin in the current quarter by 48%. This effect is larger than the respective one for the extensive margin, consistent with our hypothesis. The lagged trend for CUCMECU is 1.5% annually, or 25% over 15 years; the combined effect is 73%, which is slightly higher than the 58% estimated using the 5-year differenced RGFD results in Table 2. The 15-year lead trend for CUCMECU decreases the intensive margin by 0.5% annually, or 7% over 15 years.

Similar results are found for FTA and are reported in Table 3.

Finally, how do we reconcile the declining 15-year trend effect for annual leads on trade flows using annual data with the earlier 5-year differenced RGFD specifications that found positive 5-year lead effects? These are reconciled readily by examining Figs. 1 and 2. Fig. 1 reports the annual estimated effects of a CUCMECU using the 15-year lead and lag linear trends. 15–20 years before date of entry, country pairs that form a CUCMECU tend to trade more than country pairs that do not form an agreement; this is the selection effect. However, over the 15 years prior to date of entry, this selection effect declines; this causes the lead trend effect – which is small relative to current and lagged effects – to virtually disappear. Following date of entry, the CUCMECU increases trade, and the effect grows over time. Importantly, note that upon the date of entry the intensive margin effect (0.393) dominates the extensive margin effect (0.06). Yet, after 15 years, the intensive margin effect is a much smaller share of the total trade flow effect. After 15 years, the extensive margin effect is three times the current-period extensive margin effect. Fig. 2 provides similar results using the 15-year lead and lag linear trends for FTAs. The notable difference is that the lagged effects on the extensive margin are relatively larger for FTAs than for CUCMECUs. Thus, the annual data provide evidence that there is a material selection effect 15 years prior to formation of CUCMECUs and FTAs. However, the effect dissipates and the annual data provide convincing supporting evidence of robust partial effects

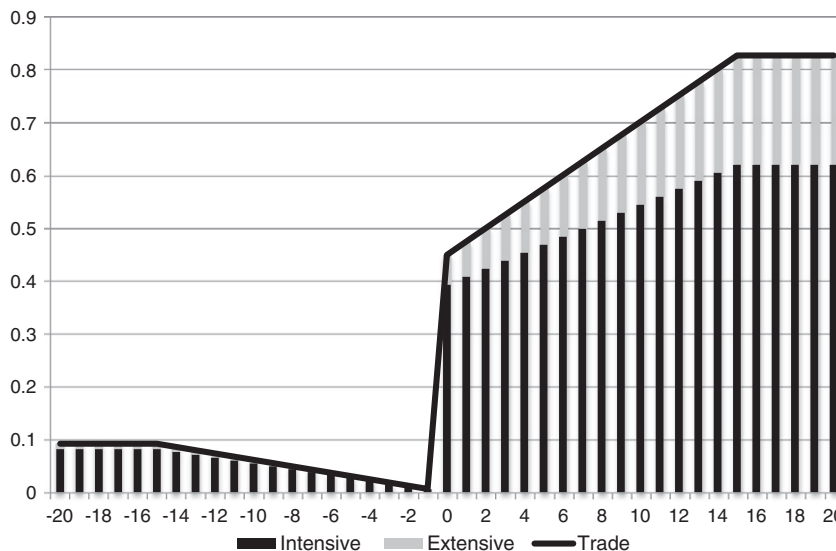


Fig. 1. CUCMECUs.

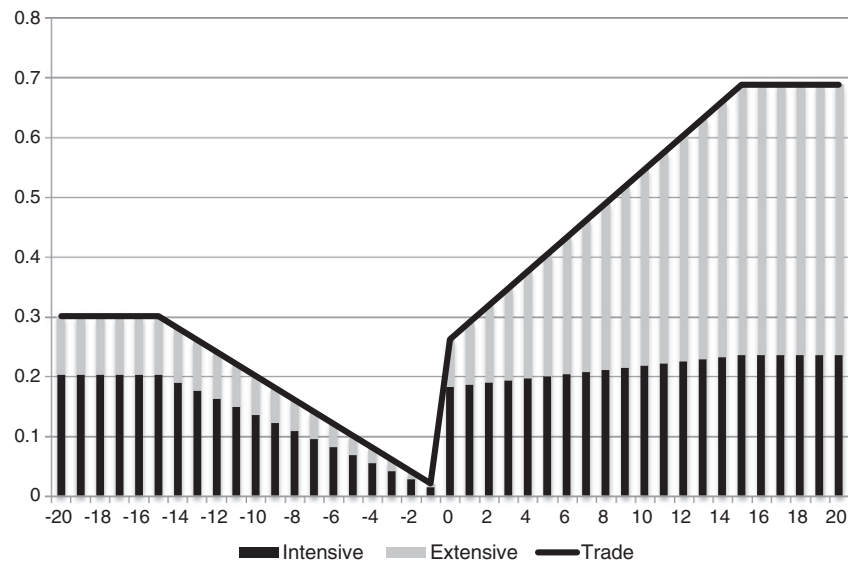


Fig. 2. FTAs.

of CUCMECUs and FTAs on aggregate trade flows, extensive margins, and intensive margins that take 10–15 years following date of entry.³²

5. Conclusions

Recent developments in estimating ex post the average treatment effects (ATEs) of economic integration agreements (EIAs) on international trade flows to account for the endogeneity of such agreements have led to larger and more precise estimates than found earlier in the gravity-equation literature. Separately, recent developments in international trade theory allowing for firm heterogeneity and fixed export costs have allowed researchers to explore the effects of trade-policy liberalizations on various extensive and intensive “margins of trade,” but most have used simulations. Only one study has explored econometrically the effect of EIAs on the intensive and extensive goods margin of trade with a large number of country-pairs’ trade flows, product categories, years, and EIAs, and no study has explored the relative effects on the margins of trade of alternative types of EIAs – partial trade agreements, free trade agreements, customs unions, etc. – much less the relative timing of such effects.

Extending methodology established in Hummels and Klenow (2005) and Baier and Bergstrand (2007), this paper provided the first evidence using gravity equations of both the intensive and extensive (goods) margins being affected by economic integration agreements (EIAs) employing a panel data set with a large number of country pairs, product categories, and EIAs from 1962 to 2000. We found the first evidence of the differential partial effects of various “types” of EIAs on these intensive and extensive margins of trade with deeper integration agreements having larger impacts on aggregate trade flows, extensive margins, and intensive margins than shallower agreements. We found the first evidence of a novel differential “timing” of the two margins’ ATEs, with intensive margin effects occurring sooner than extensive margin effects for deep EIAs, FTAs, and one-way PTAs, consistent with two recent theoretical studies. Finally, we found relative sizes of (cumulative) intensive margin and extensive margin effects consistent with implications of a gravity equation based upon a standard Melitz model. These results were robust to correcting for potential biases from sample selection, firm heterogeneity, and reverse causality.

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Appendix A. Supplementary data

Supplementary data to this article can be found online at <http://dx.doi.org/10.1016/j.jinteco.2014.03.005>.

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³² Similar findings were obtained using quadratic trends.

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