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Economic Determinants of Free Trade Agreements Revisited: Distinguishing Sources of Interdependence

*Scott L. Baier, Jeffrey H. Bergstrand, and Ronald Mariutto**

Abstract

One of the most notable international economic events since 1990 has been the enormous increase in the number of free trade agreements (FTAs). While Baier and Bergstrand were the first to show empirically the impact of a country-pair's economic characteristics on the likelihood of the pair having an FTA, the literature has been extended to demonstrate the importance empirically of FTA "interdependence"—the effect of other FTAs on the probability of a pair having an FTA. In the context of the Baier–Bergstrand framework, this paper delves deeper into the sources of interdependence—an "own-FTA" effect and a "cross-FTA" effect. The authors argue that the own-FTA effect (the impact on the net welfare gains of an FTA between two countries owing to either already having other FTAs) likely dwarfs the cross-FTA effect (the impact on the net welfare gains of an FTA between the pair owing to other FTAs existing in the rest of the world, or ROW). Augmenting a parsimonious logit model with simple "multilateral FTA" and "ROW FTA" terms to differentiate the own and cross effects empirically, it is shown that the marginal impact on the probability of a country-pair having an agreement of either country having one more FTA with a third country is 50 times that of one more FTA between another pair in ROW. The results suggest that "domino (own-FTA) effects" have far exceeded "competitive liberalization (cross-FTA) effects" in the proliferation of FTAs.

1. Introduction

One of the most notable economic events since 1990 has been the large increase in the number of bilateral and regional free trade agreements (FTAs) in existence from year to year.[1] In this period, international trade economists have mostly debated related normative questions—such as whether such agreements are on net welfare increasing or decreasing for member countries and/or for nonmembers—and related positive questions—such as whether preferential agreements are "stumbling" or "building" blocks toward global free trade. However, the profession has begun to provide empirical models that actually explain which pairs of countries have FTAs in a given year, starting with Baier and Bergstrand (2004).

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* Bergstrand: Department of Finance, Mendoza College of Business, and Kellogg Institute for International Studies, University of Notre Dame, Notre Dame, IN 46556 USA and CESifo, Munich, Germany. E-mail: bergstrand.1@nd.edu. Baier: John E. Walker Department of Economics, Clemson University, Clemson, SC 29634 USA. Mariutto: ConAgra, 1 Conagra Drive, Omaha, NE 68102 USA. Baier and Bergstrand are grateful to the National Science Foundation for financial support under grants SES-0351018 (Baier) and SES-0351154 (Bergstrand). We are also grateful for comments from an anonymous referee and participants at presentations at the Midwest International Economics Meetings, University of Notre Dame, and ETH University in Zurich. We thank Adam Ayers, R. Ethan Braden, Francisco Briseno, Emma Buckley, Brandon Caruthers, Tom Foote, Mitch Gainer, Pedro Gimenez, Nick Goode, Drew Hill, Christine Hsieh, Tahir Imtiaz, Humberto Kravetz, Cherrica Li, Lindsey Lim, Ben O'Neill, Mo Sabet, James Schappler, Andrew Weiler, Alex Wheeler, and Chris Wittman for excellent research assistance on construction of the EIA data base.

Baier and Bergstrand (2004), or BB, used a numerical version of a Krugman-type general equilibrium monopolistic competition model of international trade to show that the net economic welfare gains for two countries of having an FTA in a given year were related positively to the two countries' economic sizes (gross domestic products, or GDPs), similarity of GDPs, proximity to each other, joint remoteness from the rest of the world (ROW) and relative capital–labor ratios (up to a point). Motivated by comparative statics from this six-country theoretical model, BB employed a qualitative choice model to explain the likelihood of country-pairs having FTAs using these variables. The model explained 73% of the cross-sectional variation for 1996 among 1431 pairings of 54 countries, all RHS variables had the expected coefficient signs as suggested by theory and 85% (97%) of the country-pairs with FTAs (without FTAs) were predicted correctly.

However, BB did not address systematically the influence on the likelihood of a particular country-pair *ij* forming an FTA of *i*'s or *j*'s existing FTAs with "third countries" k ($k \neq i, j$) *or* the influence on this likelihood of existing FTAs among other "country-pairs" *kl* in the ROW ($k, l \neq i, j$).[2] The purpose of this paper is to delve deeper into the sources of FTA "interdependence" and distinguish empirically two such sources for helping to explain the proliferation of FTAs—which we will term "own-FTA" and "cross-FTA" effects—and in the context of the BB theoretical framework. The own-FTA effect refers to the impact on the net welfare gains of an FTA between two countries owing to either already having other FTAs ("third country" effects). For example, this would be the effect of an FTA between Canada and the USA on the likelihood of an FTA forming between Mexico and the USA. The cross-FTA effect refers to the impact on the net welfare gains of an FTA between the pair owing to other FTAs existing in the ROW ("third country-pair" effects). For example, this would be the effect of an agreement between France and Germany on the likelihood of an FTA forming between Canada and the USA; observers in the 1980s questioned whether the growing European Community fostered the formation of the Canadian–USA FTA in 1989. While this is not the first paper to address FTA interdependence empirically, it is the first to distinguish empirically between these two complementary sources of interdependence simultaneously—and, in particular, their relative quantitative importance. The issue is important because early arguments about "domino (own-FTA) effects" (cf. Baldwin, 1993, 1995) concerned whether an FTA formation between a pair of countries was influenced by one of the countries already having an FTA with a "third country" (e.g. Economic Community expansions), whereas early arguments about "competitive liberalization (cross-FTA effects)" (cf. Bergsten, 1996) concerned whether an FTA formation between a pair of countries was influenced by formations of FTAs between other "country-pairs" in the ROW (e.g. Economic Community's expansion influencing the formation of North Atlantic Free Trade Agreement (NAFTA)).

While the notions of FTA domino effects, competitive liberalization, contagion and interdependence have existed since 1993, only three papers have attempted to quantify its importance. However, none of the three has addressed precisely the relative quantification of third country (domino) effects from third country-pair (competitive liberalization) effects, which is our goal. The first paper to address empirically the influence of FTA interdependence on the likelihood of countries i and j having an FTA (FTA_{ij}) in a subsequent year is Egger and Larch (2008), or EL. Motivated by Baldwin's domino theory of potential trade diversion of nonmembers, EL argued that the existence of an FTA between countries k and l (a "third country-pair") would increase the likelihood of FTA_{ij} (either by joining an existing one or forming a new one), with the effect decreasing in the bilateral distance between country-pairs *ij* and *kl*. To capture interdependence empirically, EL used spatial econometrics to implement a "spatial lag" for every pair *ij*, a function of all third-country-pairs *kl* ($kl \neq ij$), but allowing $kl = il$ and $kl = kj$ to accommodate Baldwin's domino effects.[3] While EL distinguished empirically between the effects of the spatial lag on enlargements versus new FTAs, the inclusion of only a single "aggregate" spatial lag precluded distinguishing empirically own-FTA from cross-FTA effects of interdependence. The goal of this paper is to distinguish empirically these two effects as well as introduce a simpler technique.

Motivated also by the domino effect of potential trade diversion of nonmembers, Baldwin and Jaimovich (2012), or BJ, similarly employ a "spatial lag" to capture interdependence effects (termed domino or "contagion" in their paper). Unlike EL, BJ's spatial lag was constructed to capture only the effects on the likelihood of FTA_{ij} of *i*'s and *j*'s existing FTAs with "third countries" *k*, that is, domino effects. Both papers found an economically and statistically significant effect of their aggregate spatial lag on the likelihood of a country-pair *ij* having an FTA in a later period (5-years later in EL; 1-year later in BJ), confirming the presence of interdependence (in EL's terms) or contagion (in BJ's terms), respectively. In a related third paper, Chen and Joshi (2010), or CJ, include two dummy variables, one to capture the effects on the probability of FTA_{ij} of either *i* or *j* having an existing FTA with any "third country" *k* (one or more FTAs) and one to capture the effects of both *i* and *j* having an existing FTA with the same third-country *k*. However, in the context of a three-country model, CJ only address own-FTA effects, precluding the effects of FTAs of third-country-pairs *kl* (k, $l \neq i, j$) on the likelihood of FTA_{ij}. Thus, EL was the first to show empirically that FTA interdependence matters, but could not distinguish with their single aggregate spatial lag own-FTA from cross-FTA effects. By contrast, BJ and CJ found evidence of own-FTA effects, but ignored cross-FTA effects.[4]

This paper offers three potential contributions to this literature. First, we will argue that the own-FTA effect likely dwarfs quantitatively the cross-FTA effect for two reasons. Own-FTA (cross-FTA) effects reflect positive (negative) terms-of-trade effects that tend to increase (decrease) the net utility gains from FTA_{ij}. Also, own-FTA effects include additionally a role for "tariff-complementarity," which cross-FTA effects do not have, which increase the net utility gains from FTA_{ij}.[5]

Second, we formulate and estimate a simple BB logit (and, in a robustness analysis, probit) equation predicting the probability of two countries having an FTA as a function of both countries' GDP sizes, GDP similarities, bilateral distance, remoteness—factors addressed in BB—and additionally "multilateral and ROW FTA indexes" to capture the own-FTA and cross-FTA effects, respectively, without having to employ the more demanding spatial econometrics used in EL and BJ.[6] In particular, our approach can distinguish empirically the effects of the number of a country-pair's own FTAs with other countries from the number of (third-country-pair) cross-FTA effects on the likelihood of a pair forming an agreement, which the single spatial lags used in EL and BJ and the dummy variables in CJ did not. Moreover, our approach is much simpler than using spatial econometrics.[7] Our logit model generates two interesting empirical findings. We find that the marginal own effect on the probability of FTA_{ij} of either *i* or *j* having an FTA with a third country *k* far exceeds the marginal cross effect on the same probability of countries *k* and *l* (k, $l \neq i$, j) having an FTA—consistent with our theoretical conjecture. The own-FTA effect on the probability of FTA_{ij} is approximately 50 times that of the cross-FTA effect. Moreover, our logit model has a pseudo-R^2 of 56% compared with only 2–33% in BJ and CJ in comparable specifications (without fixed effects) and a pseudo-R^2 of 80% compared with only 24–60% in comparable specifications (with fixed effects) in EL, BJ and CJ.

Third, using our panel of pairings of 146 countries for 46 years (with over 350,000 observations), we employ a "sensitivity–specificity" analysis to establish the optimum cutoff probability for whether or not—according to the model's predictions—a country-pair should have a bilateral FTA formed in a given 5-year period. Based on this, we predict correctly 91% of the actual FTA formations (enlargements) for every 5-year-period from 1960 to 2005 and predict correctly also 91% of the time "No-FTAs" when no FTAs existed for the same periods excluding fixed effects in our model.[8] Moreover, if we raise the rate of "true negatives" (or No-FTAs) to 97% as in BB, which increases the cutoff probability, the "true positives" rate falls only to 75%, almost as high as that in BB for only a single cross-section of 1431 pairings among 53 countries in 1996 (which was 85%). Also, we find that the percentage correctly predicted tends to rise (fall) when the multilateral and ROW FTA indexes are included (excluded). The results confirm that competitive liberalization arising from other third-country-pair FTAs has been a force behind the increase from year to year in the number of FTAs, but the own-FTA effect is likely to have been a much more important force behind this increase over time.

The remainder of this paper is as follows. In section 2, we discuss theoretical considerations from BB for motivating our econometric model. In section 3, we provide the econometric specification and data. In section 4, we discuss the main empirical results and provide a robustness analysis. In section 5, we discuss the ability of the model to predict FTAs correctly. Section 6 concludes.

2. Theoretical Considerations

Our starting point is the general equilibrium Krugman-type monopolistic competition model of international trade in BB. The model featured a six-country world with three continents; inter-continental trade costs could range from 0 to infinity. Also, each country had intra-continental trade costs ranging from 0 to infinity. Consumers had Cobb–Douglas utility between the goods in two industries and constant elasticity of substitution (CES) utility for the differentiated varieties in each industry. Each good was produced using capital and labor, with some of each factor used for fixed setup costs. Firms were homogeneous in each monopolistically competitive industry and all factors in each country were full employed. The government of each country was assumed to maximize national welfare.

A numerical version of the model showed theoretically that two (of the six) countries i and j would have larger net utility gains from an FTA the larger their economic (GDP) sizes, the more similar their GDPs, the closer the two countries to each other, the more remote the two countries from ROW, the larger their relative factor endowment differences (up to a point) and the smaller their relative factor endowment differences relative to the ROW's. However, BB did not examine the effects of existing FTAs on the welfare gains of subsequent FTAs.

In the context of the model in BB, we consider here two further hypotheses. The two hypotheses are distinguished because Hypothesis 1 addresses cross-FTA effects and Hypothesis 2 addresses own-FTA effects. Figures 1 and 2 illustrate the two hypotheses, Hypothesis 1 and Hypothesis 2, respectively. Figure 1 illustrates the case of two countries, $1A$ and $1B$—say, the USA and Canada—forming an FTA conditioned upon two other countries, $2A$ and $2B$—say, France and Germany—already having an FTA. By contrast, Figure 2 illustrates the case of two countries, $1A$ and

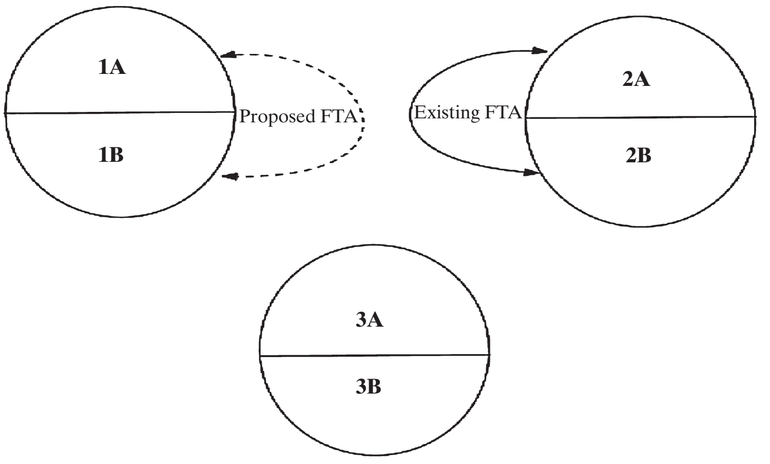

Figure 1. Hypothesis 1 (Cross-FTA Effect)

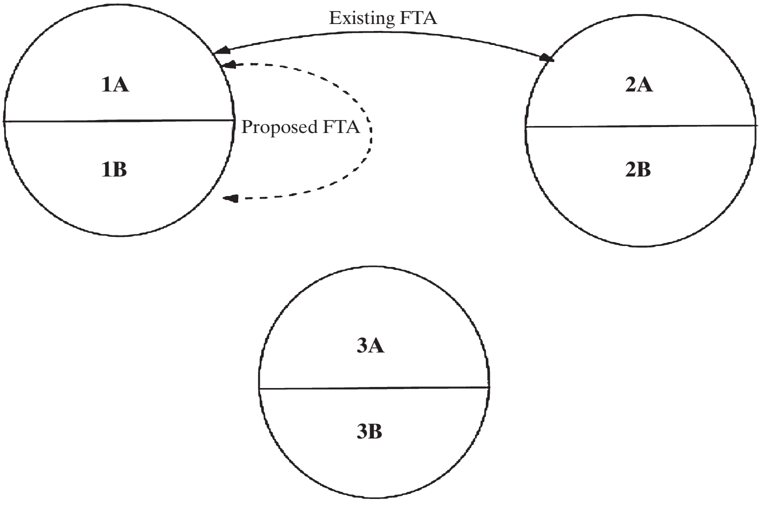

Figure 2. Hypothesis 2 (Own-FTA Effect)

1*B*—say, the USA and Mexico—forming an FTA conditioned upon one of the countries, 1*A* (say, the USA), already having an agreement with another country, 2*A* (say, Canada). It is important to note that—while countries 2*A* and 2*B* represent two countries on different continents—the BB framework allows inter- and intra-continental transport costs to vary between zero and prohibitive. Moreover, most models discussing potential trade diversion, terms-of-trade impacts and tariff-complementarity effects omit natural trade costs.

In the following, we will discuss economically the two hypotheses in the context of BB and will refer to the model and figures in that paper. We will provide also numerical comparative static effects of cross-FTAs and own-FTAs in the context of the BB model in an Online Appendix (see Supporting Information at the end of the paper). The numerical comparative statics in this Appendix are based upon the exact same theoretical and calibrated numerical model as in BB, with the exception that there is only one industry; see BB for details on the underlying theoretical model with two industries.[9]

Hypothesis 1 (Cross-FTA). *The utility gain from an FTA between two countries 1A and 1B increases owing to an existing FTA between two other countries (2A, 2B)—on the same or different continents—as a result of potential trade diversion, trade creation and terms-of-trade effects.*

In the context of BB, consider the case of two countries 1*A* and 1*B* forming a bilateral FTA. Conditioned upon no other FTAs in existence, such an FTA tends to improve welfare more the higher the level of initial tariffs. With high initial tariffs, the welfare gain from creating more trade between 1*A* and 1*B* dominates the welfare loss from trade diversion with respect to nonmember countries. Moreover, terms-of-trade (real income) improvements for the members are larger the higher the initial tariffs, adding to the welfare gains. The larger the elasticity of substitution among varieties, the larger tend to be the welfare gains also for 1*A* and 1*B*; if products have low substitutability, trade diversion of valued differentiated products from nonmembers reduces on net the welfare gains.

Suppose now that countries 2*A* and 2*B* already have an FTA. Is the welfare gain from $FTA_{1A,1B}$ now larger or smaller, given the existence of $FTA_{2A,2B}$, and all else constant? Intuitively, when 2*A* and 2*B* form an FTA, each of 1*A* and 1*B* experiences trade diversion, a loss of terms of trade and erosion in real income. When a country pair (2*A*, 2*B*) is remote, there are negligible volume-of-trade and terms-of-trade (real income) effects on 1*A*'s utility from the formation of $FTA_{2A,2B}$ because there is little trade to be diverted between country 1*A* and countries 2*A* and 2*B*. However, if inter- (and intra-) continental trade costs are low, then 1*A* trades considerably with 2*A* and 2*B*. An FTA between 2*A* and 2*B* causes substantive trade diversion for 1*A*, eroding 1*A*'s volume of trade with 2*A* and 2*B* and 1*A*'s utility and real income, but improving 1*A*'s volume of trade with 1*B*. Consequently, the formation of $FTA_{1A,1B}$ has an even larger impact on 1*A*'s utility—in the presence of $FTA_{2A,2B}$ than in its absence—because the elimination of tariffs from $FTA_{1A,1B}$ on the greater volume of trade between 1*A* and 1*B* owing to $FTA_{2A,2B}$ more than offsets the terms-of-trade loss caused by trade diversion from $FTA_{2A,2B}$. $FTA_{2A,2B}$ effectively has made countries 1*A* and 1*B* more "economically remote" and this isolation has made 1*A* and 1*B* economically more natural trade partners, enhancing the gains from an FTA. Referring to either Figure 1 or Figure 2 in BB, we know that increased "remoteness" of a country-pair tends to enhance the welfare gains from an FTA.[10] Consequently, country 1*A*'s (and, by symmetry, 1*B*'s) "demand for membership" in an FTA with country 1*B* (1*A*) will tend to be higher if 2*A* and 2*B* have an existing FTA. The positive difference is the role of "third-country-pairs" creating competitive liberalization.

Hypothesis 2 (Own-FTA). *The utility gain from an FTA between two countries 1A and 1B increases owing to the existence of an FTA between either of these countries with another (third) country and the gain is likely larger than in the previous case.*

Consider again the case of two countries 1*A* and 1*B* forming a bilateral FTA. Suppose now instead that 1*A* and 2*A* already have an FTA. Country 1*A*'s "demand for membership" in an FTA with country 1*B* will tend to increase if 1*A* has an existing FTA with another country (say, 2*A*). Moreover, the effect is largest when trade costs are low. Note importantly that this is not due to potential trade diversion of 1*B*; country

$1A$ is already in an agreement with $2A$, so this is different from the trade diversion arguments in EL and BJ. The economic intuition behind this is the following. At high trade costs, there is little trade between $1A$ and $2A$ so there can be little impact of $FTA_{1A,2A}$ on the gains to $1A$ from $FTA_{1A,1B}$. However, at low transport costs, $1A$ trades considerably with $2A$, and $FTA_{1A,2A}$ causes considerable trade diversion for $1A$ with $1B$, unlike the case of $FTA_{2A,2B}$ which increases $1A$'s trade with $1B$.

Two implications are worth noting. First, in contrast with Hypothesis 1, since $1A$ and $1B$ are trading less in the presence of $FTA_{1A,2A}$ than in its absence, this lower volume of trade erodes the relative gain to $1A$'s welfare of $FTA_{1A,1B}$. Second, one cannot ignore that $FTA_{1A,2A}$ increased the terms of trade and real income of country $1A$ (as well as that of $2A$). Referring to Figure 2 in BB, increased real income tends to enhance a country's welfare gain from an FTA. This improvement in terms of trade and real income has a positive benefit for improving $1A$'s utility gain from $FTA_{1A,1B}$, conditioned upon $FTA_{1A,2A}$. The combination of these effects suggests that $1A$ has an incentive to form an FTA with $1B$; this effect is analogous to the notion of "tariff-complementarity" addressed in Bagwell and Staiger (1998).[11]

We emphasize the relatively larger potential benefits from $FTA_{1A,1B}$ from the existence of $FTA_{1A,2A}$ compared with the existence of $FTA_{2A,2B}$ (as measured by the percentage change in utility). This is because $FTA_{1A,2A}$ causes a large increase in terms of trade and real income for $1A$ while $FTA_{2A,2B}$ causes a loss of terms of trade and real income for $1A$, even though $FTA_{1A,2A}$ leads to less trade volume between $1A$ and $1B$ and $FTA_{2A,2B}$ leads to more trade volume between $1A$ and $1B$. Hence, the percentage gain in utility for $1A$ from $FTA_{1A,1B}$ conditioned on $FTA_{1A,2A}$ is greater than that from $FTA_{1A,1B}$ conditioned on $FTA_{2A,2B}$ owing to the terms-of-trade effects. Moreover, while $2A$ experiences some trade diversion with respect to $1B$ as a result of $FTA_{1A,1B}$, $2A$ still has an incentive to be in an FTA with $1A$.[12]

Finally, as we would expect based upon the "domino effect" hypothesis, $1B$ suffers trade diversion and loss of real income from $FTA_{1A,2A}$. Despite the loss of real income, $1B$ on net benefits from an FTA with $1A$, raising $1B$'s demand for membership in $FTA_{1A,1B}$. Consequently, these considerations suggest that an increase in the number of FTAs that, say, country $1A$ has with other (non-$1B$) countries increases the net utility gains of $FTA_{1A,1B}$. These considerations also suggest that the marginal own-FTA effects on the probability of $FTA_{1A,1B}$ will exceed the marginal cross-FTA effects on this probability.

3. Econometric Issues and Data

Econometric Issues

The econometric framework employed is the qualitative choice model of McFadden (1975, 1976), as in BB. A qualitative choice model can be derived from an underlying latent variable model. For instance, let y_{ijt}^* denote an unobserved (or latent) variable. As in Wooldridge (2000), let y_{ijt}^* in the present context represent the percentage difference in utility levels from an action (formation of an FTA) between countries i and j in year t, where:

$$y_{ijt}^* = \alpha + x_{ijt}\beta + \varepsilon_{ijt} \tag{1}$$

where α is a parameter, x_{ijt} is a vector of explanatory variables (i.e. economic characteristics), β is a vector of parameters and error term ε_{ijt} is assumed to be independent of x_{ijt} and to have a logistic distribution; we will also consider in the sensitivity analysis the standard normal distribution for ε_{ijt}. In the context of the BB model formally, $y^{*}_{ijt} = \min(\Delta U_{it}, \Delta U_{jt})$ where ΔU_{it} (ΔU_{jt}) denotes the percentage change in utility for the representative consumer in i (j) in year t. Hence, both countries' consumers need to benefit from an FTA for their governments to form one, as in BB.

Since y^{*}_{ijt} is unobservable, following BB we define an indicator variable, FTA_{ijt}, which assumes the value 1 if two countries have an FTA and 0 otherwise, with the response probability, Pr, for FTA:

$$\Pr(FTA_{ijt} = 1) = \Pr(y^{*}_{ijt} > 0) = G(\mathbf{x}_{ijt}\beta), \tag{2}$$

where $G(\bullet)$ is the logistic cumulative distribution function, ensuring that $\Pr(FTA_{ijt} = 1)$ is between 0 and 1.[13] While the statistical significance of the logit estimates can be determined using t-statistics, the coefficient estimates can only reveal the sign of the partial effects of changes in x on the probability of an FTA, owing to the nonlinear nature of $G(\bullet)$. Drawing upon analogy to the labor literature, we assume the existence of a "reservation cost" to forming an FTA (denoted y^{*R}_{ijt}). Hence, the gain in utility from forming/joining an FTA must exceed this cost (e.g. political and/or administrative cost of action) in order for an FTA "event" to occur. If $y^{*}_{ijt} - y^{*R}_{ijt} > 0$, then the FTA event for the pair of countries occurs at time t. Initially in our empirical specifications, we assume y^{*R}_{ijt} is exogenous and constant; however, y^{*R}_{ijt} may differ across country-pairs and also such costs may vary over time. However, in our sensitivity analysis, we can account for the possibility that the reservation costs differ across county pairs; this is easily accounted for using country-pair fixed effects (in logit specifications). Moreover, such reservation costs may be influenced by trends in policy toward "globalization." We will also be able to account for such trends over time by including a set of time dummies. All of this is explored in the empirical sensitivity analysis.

Intuition for Multilateral FTA Terms

The theoretical considerations above, alongside earlier considerations discussed in BB, suggest that x_{ijt} should be influenced not only by the distance between countries i and j, their remoteness, and the economic size and similarity of countries i and j but also by "multilateral" indexes of each of i's and j's other FTAs (for own-FTA effects) and an index of all FTAs other than those with i or j (for cross-FTA effects). While measurement of distance, economic size and economic similarity is straightforward, indexes of "multilateral FTAs" for i and j and an index of all other non-ij FTAs—henceforth, for tractability, termed ij's "ROW FTAs" index—are not readily observed.

However, intuition for the construction of our multilateral and ROW FTA indexes becomes transparent once we re-emphasize one of our main goals: to estimate the marginal impacts on the probability of FTA_{ij} of either country i or j having one more existing FTA with a third country k and of there existing one more FTA in the ROW (say, between k and l). Thus, for empirical purposes the appropriate measure of the multilateral FTA index for, say, country i is simply the count of i's FTAs with other (non-j) countries. Analogously, the appropriate measure of the multilateral FTA index for j is the count of j FTAs with other (non-i) countries. The appropriate measure of the ROW FTA index for pair ij is then simply the count of all FTAs in the world that exclude i and j.[14]

Based on these considerations, we define a multilateral index of country *i*'s FTAs with every other (non-*j*) country lagged 5 years (to avoid endogeneity bias), $MFTA_{i,t-5}$, which is an unweighted sum of country *i*'s indexes of FTAs with all other countries (excluding *j*):

$$MFTA_{i,t-5} = \sum_{k \neq j}^{N} FTA_{ik,t-5} \tag{3}$$

where $FTA_{ik,t-5}$ is a binary variable assuming the value 1 if *i* and *k* have an FTA in year $t-5$, and 0 otherwise. Analogously, we define for *j*:

$$MFTA_{j,t-5} = \sum_{k \neq i}^{N} FTA_{jk,t-5}. \tag{4}$$

It follows that we can define the cross-FTA index for country-pair *ij*, $ROWFTA_{ij,t-5}$, as:

$$ROWFTA_{ij,t-5} = \sum_{k \neq i,j}^{N} \sum_{l \neq i,j}^{N} FTA_{kl,t-5}. \tag{5}$$

Hypothesis 1 ("cross-FTA" effect) suggests that the coefficient estimate for $ROWFTA_{ij,t-5}$ should be positive. Hypothesis 2 suggests that the coefficient estimates for $MFTA_{i,t-5}$ and $MFTA_{j,t-5}$ should be positive and their marginal response probabilities larger than those for $ROWFTA_{ijt}$.[15]

An alternative measure might recognize that each bilateral FTA component of these indexes should be weighted by its relative economic importance. This suggested alternative GDP-weighted multilateral and ROW indexes:

$$MFTAY_{i,t-5} = \sum_{k \neq j}^{N} Y_{k,t-5} FTA_{ik,t-5} \tag{6}$$

where $Y_{k,t-5}$ is country *k*'s GDP in year $t-5$. We define $MFTAY_{j,t-5}$ and $ROWFTAY_{ij,t-5}$ analogously. We apply this alternative weighting method in the sensitivity analysis.[16]

Finally, alternative weights that come to mind are bilateral-trade-share weights or factors that might influence bilateral trade shares, such as inverse-bilateral-distances or GDPs divided by bilateral distances. BJ used bilateral trade shares, as did EL in a sensitivity analysis of their spatial-lag construction. However, as both studies noted, such shares may create an endogeneity bias. Consequently, EL relied upon inverse-distance weights in their construction of their primary spatial lags. Scaling by inverse distances may create problems because of the roles of inter- and intra-continental transport costs. We examine the sensitivity of the results to the roles of inter- and intra-continental transport costs later when we estimate the marginal response probabilities separately for trading partners on the same or on different continents.

Multilateral Resistance and Other Data Issues

Since Tinbergen (1962), gravity-equation analyses of bilateral trade flows have measured the presence or absence of an FTA between a country-pair using a binary variable. Following those studies and BB, variable FTA_{ijt} will have the value 1 for a pair of countries (*i*, *j*) with an FTA (specifically, FTA, customs union, common market, or economic union) in year *t*, and 0 otherwise; we exclude one-way and two-way

Table 1. Data Description[a]

Integration index	*Count*	*Percent of total*	*Percent of subtotal*
0 (None)	336,640	69.1	85.8
1 (1-way PTA)	33,821	7.0	8.6
2 (2-way PTA)	11,035	2.3	2.8
3 (FTA)	7,498	1.5	1.9
4 (Customs Union)	1,547	0.3	0.4
5 (Common Market)	1,085	0.2	0.3
6 (Economic Union)	643	0.1	0.2
Subtotal	392,269	—	100.0
Missing observations	94,641	19.5	
Total	486,910	100.0	

[a] Total observations are based upon 146 countries ($146 \times 145/2 = 10{,}585$ pairings) for 46 years (1960–2005). Missing data refers to country pairs where in a given year one of two countries did not officially exist. See data source at www.nd.edu/~jbergstr.

"preferential" trade agreements (where "preferential" denotes only partial liberalization, not "free" trade). This variable was constructed using all bilateral pairings among 195 countries in the world annually from 1960 to 2005.[17] A decomposition of cells is provided in Table 1.

The only other data needed are real GDPs, bilateral distances, a dummy variable assuming the value 1 (0) if two countries are on the same continent ($CONT_{ij}$), and indexes of "remoteness." In order to employ a consistent real GDP data set for such a long period, we use real GDP data from Maddison (2009). However, the cost of a consistent real GDP panel data set for such a long time period is number of usable countries. This lowers the number of countries from 195 to 146. We construct for every country-pair the variable $SUMGDP_{ij,t-5}$, which is the natural log of the sum of i's and j's real GDPs 5 years prior to year t. We measure the dissimilarity of economic sizes using $DIFGDP_{ij,t-5}$, which is the absolute value of the difference in the log of each country's real GDP. Bilateral distances are calculated from great-circle distances using latitudes and longitudes between economic centers from the CIA's *World Factbook*, as is standard. $DIST_{ij}$ refers to the natural logarithm of the bilateral distance between the two countries i and j. However, measures of "remoteness" of a country-pair are not readily observable. We now address this issue briefly.

Recent studies by EL, BJ and CJ have followed BB and used a simple average of the logarithms of the simple averages of each of countries i's and j's bilateral distances to all other countries to measure a pair of countries' "remoteness," cf. BB (2004, p. 40). This variable typically has a positive coefficient estimate sign and statistical significance. However, there is no explicit theoretical foundation for its formulation. It turns out that a formulation very close to this surfaces from recent developments in the theoretical foundations for the gravity equation. These recent developments—based upon Anderson and van Wincoop (2003), as modified using a Taylor-series expansion in Baier and Bergstrand (2009)—provide guidance for measuring remoteness using "multilateral resistance" indexes that are very similar to our *MFTA* and *ROWFTA* indexes. For instance, for country i's multilateral resistance index for the log of distance we use either:

$$MDIST_i = \frac{1}{N}\sum_{k}^{N} DIST_{ik} \tag{7}$$

or

$$MDISTY_{it} = \sum_{k}^{N} \theta_{kt} DIST_{ik} \tag{8}$$

where $\theta_{kt} = Y_{kt}/Y_t^W$. Analogous terms apply for $MDIST_j$ and $MDISTY_{jt}$.[18] Similarly, for country i's multilateral resistance index for the binary variable $CONT_{ij}$ we define:

$$MCONT_i = \frac{1}{N}\sum_{k}^{N} CONT_{ik} \tag{9}$$

or

$$MCONTY_{it} = \sum_{k}^{N} \theta_{kt} CONT_{ik} \tag{10}$$

and the analogous terms for $MCONT_j$ and $MCONTY_{jt}$. For parsimony, in the spirit of Baier and Bergstrand (2009), we condense these multilateral resistance terms into two variables for each country-pair. For constructing the "multilateral resistance" term for distance for the unweighted case, we have:

$$MDIST_{ij} = \frac{1}{2N}\left(\sum_{k=1}^{N} DIST_{ik} + \sum_{k=1}^{N} DIST_{jk}\right) \tag{11}$$

and analogously for the GDP-share-weighted case. For $CONT$, we use:

$$MCONT_{ij} = \frac{1}{2N}\left(\sum_{k=1}^{N} CONT_{ik} + \sum_{k=1}^{N} CONT_{jk}\right). \tag{12}$$

and analogously for the GDP-shared-weighted case (allowing for time variation in the GDP weights). See Baier and Bergstrand (2009) for details.

4. Empirical Results

Main Results

Table 2 provides the main empirical results. Specification 1 provides the results where the right-hand side variables are—in the case of unweighted averages—time-invariant variables. Specification 1 shows that $DIST_{ij}$, $CONT_{ij}$ and $MCONT_{ij}$ all have the expected signs and are statistically significant at conventional significance levels (1%). Bilateral distance has a negative effect on the probability of an FTA, while being on the same continent has a positive effect; these results are in line with the cross-sectional findings in BB and our earlier discussion for "core" economic variables. $MDIST_{ij}$ and $MCONT_{ij}$ both have negative effects on the likelihood of an FTA. While the coefficient estimate for $MCONT_{ij}$ is as expected, that for $MDIST_{ij}$ is the opposite of our expectation, since this is effectively a measure of "remoteness." However, we

Table 2. Regression Results

Variables	*Expected sign*	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*	*(8)*	*(9)*	*(10)*	*(11)*	*(12)*	*(13)*
$DIST_{ij}$	–	–1.02*	–1.30*	–1.54*	–1.24*	–0.72*	–1.59*	–1.59*	–0.74*	–0.70*	–	–	–	–
		(–59.26)	(–66.97)	(71.65)	(–46.42)	(–74.67)	(–77.35)	(–35.42)	(–86.64)	(–58.87)				
$MDIST_{ij}$	+	–2.95*	–1.39*	0.25*	–0.36*	–0.03	0.36*	0.39*	–0.44*	–0.96*	–	–	–	–
		(–75.42)	(–37.37)	(5.56)	(–6.08)	(–1.40)	(8.86)	(4.32)	(–12.03)	(–18.63)				
$CONT_{ij}$	+	1.71*	1.45*	1.65*	1.43*	0.85*	1.57*	1.54*	1.87*	1.79*	–	–	–	–
		(50.78)	(41.53)	(40.72)	(27.85)	(45.68)	(40.36)	(18.74)	(59.79)	(41.36)				
$MCONT_{ij}$	–	–10.07*	–4.11*	–0.85*	–2.40*	–0.80*	–0.48	–0.90*	–2.91*	–4.12*	–	–	–	–
		(–67.73)	(–26.70)	(–5.10)	(–11.07)	(–9.79)	(–5.93)	(–2.70)	(–19.79)	(–20.77)				
$SUMGDP_{ij,t-5}$	+	–	0.80*	0.51*	0.49*	0.24*	0.48*	0.50*	0.37*	0.39*	3.20*	3.43*	2.81*	4.86*
			(79.20)	(44.79)	(34.54)	(42.96)	(32.85)	(21.51)	(39.00)	(29.69)	(14.08)	(12.32)	(6.05)	(7.34)
$DIFGDP_{ij,t-5}$	–	–	–0.56*	–0.51*	–0.42*	–0.24*	–0.47*	–0.53*	–0.44*	–0.39*	0.40	0.16	–0.10	–1.00*
			(–50.73)	(–40.67)	(–27.18)	(–40.12)	(–37.97)	(–20.72)	(–42.65)	(26.74)	(1.66)	(0.56)	(–0.33)	(–2.41)
$MFTA_{i,t-5(i \neq j)}$	+	–	–	0.09*	0.02*	0.05*	0.06*	0.09*	0.07*	0.01*	0.35*	0.84*	0.33*	0.97*
				(45.68)	(8.47)	(45.18)	(59.76)	(23.37)	(44.87)	(5.80)	(11.20)	(10.36)	(13.64)	(13.13)
$MFTA_{j,t-5(j \neq i)}$	+	–	–	0.12*	0.05*	0.06*	0.06*	0.11*	0.09*	0.04*	0.45*	0.88*	0.36*	0.84*
				(57.17)	(19.48)	(56.13)	(61.83)	(27.78)	(55.24)	(17.14)	(11.56)	(10.90)	(13.33)	(12.41)
$ROWFTA_{ij,t-5}$	+	–	–	0.0004*	0.002*	0.0002*	0.22*	0.0004*	0.00004	0.001*	0.009*	0.01*	0.01*	0.01
				(9.71)	(26.31)	(12.45)	(31.01)	(5.32)	(0.90)	(25.56)	(22.97)	(20.86)	(22.42)	(1.03)
CONSTANT		–116.17*	–50.03*	–13.69*	–33.08*	–11.27*	–6.01*	–14.90*	–41.27*	–55.69*	–	–	–	–
		(–60.61)	(–24.99)	(–6.31)	(–11.82)	(–10.67)	(–14.13)	(–3.44)	(–21.76)	(–4.83)				
Pair Fixed Ef.		No	No	No	No	No	No	No	No	No	Yes	Yes	Yes	Yes
Yr Dummies		No	No	No	No	No	No	No	No	No	No	No	Yes	Yes
No. Obs.		358,767	358,767	358,767	352,002	358,767	358,767	77,059	358,767	352,002	358,767	352,002	358,767	352,002
No. Pos. Obs.		10,478	10,478	10,478	3,811	10,478	10,478	2,490	10,478	3,811	10,478	3,811	10,478	3,811
Random Prob.		0.029	0.029	0.029	0.011	0.029	0.029	0.032	0.029	0.011	0.029	0.011	0.029	0.011
Pseudo R^2		0.39	0.46	0.56	0.34	0.57	0.56	0.56	0.54	0.32	0.80	0.73	0.87	0.85
log-likelihood		–28,953	–25,359	–20,663	–13,860	–20,587	–20,759	–4,792	–21,832	–14,260	–2,294	–2,034	–1,525	–1,142

Note: *Denotes statistical significance in one-tail t-test (or z-test) at 1% level.

will see shortly that this unexpected negative coefficient sign is reversed in a fuller specification. The pseudo-R^2 is 0.39. Recall that in logit (or probit) regressions the coefficient signs are meaningful, but the actual values of the coefficients are not directly interpretable; however, marginal response probabilities will be calculated later, as in BB, to examine the quantitative effects of one-standard-deviation changes in RHS variables (and also, for count variables, unit changes in the right-hand side variables).

Specification 2 in Table 2 augments Specification 1 to include the (5-year-lagged) logarithm of the joint economic size of countries i and j ($SUMGDP_{ij,t-5}$) and our measure of dissimilarity of economic sizes of i and j ($DIFGDP_{ij,t-5}$). We find that country-pairs are more likely to form an FTA the larger and more similar are their GDPs, in accordance with BB and earlier discussion. The results in Specification 2 confirm using a very large pooled cross-section time-series data set the results found for a single cross-section of a smaller number of countries in BB and are consistent with the pooled cross-section time-series results in EL, BJ and CJ.[19]

We now address Hypotheses 1 and 2. Specification 3 provides the results of augmenting Specification 2 with $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$. Specification 3 is our main specification. First, all three of these variables have statistically significant positive coefficient estimates and the coefficient estimates of the other right-hand side variables retain their same signs and remain statistically significant except $MDIST_{ij}$, which reverses its sign to the expected positive one and is statistically significant. Second, the positive coefficient estimates for $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$ all confirm Hypotheses 1 and 2. Third, the pseudo-R^2 of the logit regression is 0.56, which is substantive and very close to the pseudo-R^2 found in BB for their much smaller and select cross-section sample for the year 1996. Moreover, it is larger than the pseudo-R^2 values for comparable specifications without fixed effects in BJ and CJ of 33% and 4%, respectively; EL's specifications all included time-meaned fixed effects. We note that there are 10,478 observations with FTAs ($FTA_{ij,t} = 1$) in the sample of 358,767 observations spanning 1960–2005. Finally, we note that the coefficient estimates for $MFTA_{i,t-5}$ and $MFTA_{j,t-5}$ are substantively larger than that for $ROWFTA_{ij,t-5}$. The relatively larger coefficient estimates for the former variables are seemingly consistent with the relative quantitative predictions for the relative utility gains. However, because of the nonlinearities using logit regressions, we will delay full discussion of these relative quantitative predictions until we examine more appropriately marginal response probabilities later.

Sensitivity Analysis

Specification 3 provides the main specification for predicting later the rate of "true positives" (predicting an FTA when one exists) and the rate of "true negatives" ("No-FTA" when none exists). In this sensitivity analysis, we examine the robustness of results using Specification 3 to examining "formations" of FTAs (rather than the indicator representing "existence" of an FTA in a given year), to using probit rather than logit estimation, to using alternative weights for the various multilateral and ROW index variables, to using a panel of every 5 years (rather than annual), to using instead a duration model, to the presence of country-pair fixed effects and to inclusion of time dummies in addition to country-pair fixed effects. Finally, we report marginal response probabilities for Specification 3.

First, one concern of Specification 3 is that we are examining the "existences" of FTAs in a given year rather than their "formations." $FTA_{ij,t}$ assumes the value 1 if an

FTA exists between *i* and *j* in any year *t*, and 0 otherwise. Alternatively, we would like to consider another dummy variable for the left-hand side that assumes the value 1 in a year *t* when an FTA is formed between *i* and *j* in that year and 0 otherwise.[20] Consequently, we construct a new variable, $TFTA_{ij,t}$, which assumes the value 1 if countries *i* and *j* transitioned into an FTA in year *t*, and 0 otherwise. As a result, the number of observations with FTA formations ($TFTA_{ij,t} = 1$) is 3811, approximately one-third that for $FTA_{ij,t} = 1$. Specification 4 reports the results of replacing $FTA_{ij,t}$ with the "transition-to-FTA" binary variable $TFTA_{ij,t}$. We note that the number of total observations falls from 358,767 to 352,002 as we redefine the left-hand side dummy variable to represent the change from one year to the next in the FTA "status" of the pair. Note that the coefficient estimates in Specification 4 are qualitatively identical to those in Specification 3, with the exception of $MDIST_{ij}$ which has a negative effect now. Most importantly though, all the main results hold up; $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$ all have positive and statistically significant coefficient estimates as expected. However, predicting transitions is more challenging than predicting the existence of FTAs; the pseudo-R^2 is lower at 0.34 compared with 0.56 for Specification 3, as expected.[21]

Second, both BB and EL used probit estimation rather than logit, the latter used here. Recall that our reason for using logit is that we will include country-pair fixed effects shortly to compare our results with EL and BJ, both of which included Chamberlain (1980) time-meaned fixed effects in their probits. As clarified in Wooldridge (2002, pp. 490–492), since the logistic transformation is a linear one standard fixed effects can be readily applied without restrictions; by contrast, fixed effects have more restrictions on their implementation because of the Chamberlain (1980) "incidental parameters" issue. However, it is useful to show that the results are robust to estimation using probit. Specification 5 provides the results of re-estimating the model of determinants of existence of an FTA ($FTA_{ij,t}$) using probit. The results in Specification 5 are qualitatively identical to the corresponding logit ones in Specification 3 with one exception; the coefficient estimate for $MDIST_{ij}$ reverses signs from positive (which is expected) to negative but statistically insignificant in the probit specification. The pseudo-R^2 is 0.57, virtually identical to the pseudo-R^2 of 0.56 in Specification 3. Quantitatively, with the exception of that for $MDIST_{ij}$, all the probit coefficient estimates are approximately a half of those in the logit equation. Thus, the results are largely robust to estimation using probit instead.[22]

Third, as discussed earlier, our theoretical model provides no clear guidance for weights for the multilateral FTA indexes. Following guidance from recent theoretical developments for the gravity equation in Anderson and van Wincoop (2003) as modified by Baier and Bergstrand (2009), the two weighting methods suggested are a simple average of components or a GDP-weighted average of bilateral components. We re-estimated our main logit specification (3) using GDP-weighted values: $MFTAY_{i,t-5}$, $MFTAY_{j,t-5}$, $ROWFTAY_{ij,t-5}$, $MDISTY_{ij}$ and $MCONTY_{ij}$. The results are provided in Specification 6 for the existence of FTAs.[23] The results in Specification 3 are robust to using the GDP-weighted (or GDP-share-weighted) alternative variables; all coefficient estimates are qualitatively identical to those in Specification 6.[24]

Fourth, in the specifications used so far, we use 5-year lagged values of $SUMGDP_{ij,t-5}$, $DIFGDP_{ij,t-5}$, $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$ to predict the existence of (or transition to) an FTA for a country pair within the next 5 years. However, this large window for FTA predictions may introduce an endogeneity bias. Consequently, we re-estimated the main specification using only right-hand side and left-hand side variables with a sub-sample of every 5 years. This reduced our sample size for predicting existence from 358,767 to 77,059. The results are provided in Specification 7. We see in column (7) that all of the coefficient estimates are robust to this alternative specification which retains the 5-year lag for the right-hand side variables. The results for transition-to-FTA are also robust, but omitted from the table for brevity.

Fifth, Bergstrand et al. (2009) implemented instead a "duration analysis" of the likelihood of FTA events. Like logit and probit regressions, duration models fall within the class of "limited dependent variable" models, cf. Wooldridge (2002). These models estimate the "hazard rate," which is the instantaneous probability of leaving an initial state (No-FTA) in the interval $[t, t + dt)$ given survival up until time t. We also estimated a duration model using the same variables; the results are in Specification 8 (9) for existence of (transition to) FTAs. Columns (8) and (9) indicate that our main results from Specification 3 are robust qualitatively to using a duration model rather than a simple logit (or probit) model. All coefficient estimates (except, as before, that for $MDIST_{ij}$) are correctly signed and statistically significant.

Sixth, the results may be sensitive to omitted unobserved cross-sectional heterogeneity. As is often done in gravity-equation analyses of trade flows, one introduces country-pair fixed effects to account for unobserved heterogeneity to ensure unbiased coefficient estimates. Country-pair fixed effects are likely to account for variation in pair-specific political factors that may influence FTA formations. As noted earlier, one of the advantages of logit over probit estimation (or duration analysis) is the ability to use standard fixed effects; by contrast, such effects cannot be used in probit specifications owing to the normal distribution underlying probits. However, Chamberlain (1980) offers a methodology to include time-meaned fixed effects in probits; such effects were included in both EL and BJ. Of course, the introduction of country-pair fixed effects implies removing all time-invariant variables, i.e. $DIST_{ij}$, $CONT_{ij}$, $MDIST_{ij}$ and $MCONT_{ij}$. Specification 10 reports the results of introducing country-pair fixed effects into main Specification 3. We see that the remaining time-varying variables' coefficient estimates are significant, with only the estimate for our measure of size-dissimilarity having an unexpected positive sign but statistical insignificance. All the other four variables' coefficient estimates retain the same expected positive signs as in previous regressions. When we introduce the same fixed effects into the logit regressions using "transition-to-FTA" binary variable $TFTA_{ij,t}$, shown in Specification 11, the coefficient estimates remain positive and statistically significant with the exception again of insignificance for the coefficient estimate for the GDP-size-dissimilarity variable. Importantly, the pseudo-R^2 value for our fixed effect logits for existence of FTA in Specification 10 is 80%. This is substantially higher than the pseudo-R^2 values ranging from 55 to 60% for the comparable fixed effects specifications in EL and BJ and 7 to 24% in CJ.

Seventh, while the country-pair fixed effects specifications controlled for unobservable time-invariant factors, they did not control for unobservable time-varying factors. For instance, world GDP and technology change over time. More specific to the issues at hand, global liberalization of trade under the GATT/WTO may have had an influence on the likelihood of bilateralism being captured in the remaining time-varying RHS variables $SUMGDP_{ij,t-5}$, $DIFGDP_{ij,t-5}$, $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$. Specifications 12 and 13 add to the logit fixed-effects specifications 10 and 11, respectively, time dummies. Columns (12) and (13) report the results for the existence-of-FTA and transition-to-FTA specifications, respectively. In Specification 12 the $SUMGDP_{ij,t-5}$ ($DIFGDP_{ij,t-5}$) coefficient estimate has the expected positive (negative) sign and is statistically significant (insignificant). Moreover, the coefficient estimates for $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$ are positively signed as expected and remain statistically significant. In Specification 13, the $SUMGDP_{ij,t-5}$ and $DIFGDP_{ij,t-5}$ coefficient estimates have the expected signs and both are statistically significant. The coefficient estimates for $MFTA_{i,t-5}$ and $MFTA_{j,t-5}$ remain positively signed and statistically significant; the coefficient estimate for $ROWFTA_{ij,t-5}$ is positively signed, but statistically insignificant. These results confirm the importance of existing FTAs for enhancing the likelihood of subsequent FTAs.

Table 3 reports the marginal response probabilities, calculated at the means of the levels of all variables. Importantly, we use our main specification, Specification 3, that excludes fixed effects so as not to contaminate the predictions with pair fixed effects. We follow the approach used in BB by separating the marginal response probabilities into those calculated for "natural" trading partners (i.e. pairs on the same continent) and those for "unnatural" trading partners (i.e. pairs on different continents). There are two reasons for this here. First, as in BB, it makes little economic sense to evaluate the marginal response probabilities at the "mean" of a binary variable representing the presence or absence of being on the same continent. Second, the utility gains for a country-pair from forming an FTA are likely sensitive to the level of transportation costs. One transparent method for evaluating the influence of distance on the effects of existing FTAs on the likelihood of subsequent FTAs is to evaluate marginal response probabilities separately for natural and unnatural trading partners. The format of this table is the same as in BB.

Table 3a reports the marginal response probabilities for natural trading partners. First, for ease of reference the probability of an FTA among natural partners at the mean level of all other RHS variables is 0.1031, with a 95% confidence interval of 0.0995 to 0.1068. We now consider the effect of a one standard deviation (S.D.) increase or decrease of variables. The sixth (seventh) line of Table 3a indicates that a one S.D. increase in $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) increases the probability of $FTA_{ij,t}$ to 0.1306 (0.1381). Each of these probability changes is statistically significant at the 95% level, but not significantly different from one another.[25] By contrast, a one S.D. increase in $ROWFTA_{ij,t-5}$ increases the probability of $FTA_{ij,t}$ to only 0.1101, which is also a statistically significant change. The difference in the marginal response probabilities for $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) and $ROWFTA_{ij,t-5}$ is economically and statistically significant. Moreover, the difference in such probabilities is as expected; a one S.D. change in $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) has a quantitatively larger impact on the likelihood of $FTA_{ij,t}$ than does a one S.D. change in $ROWFTA_{ij,t-5}$. In fact, a typical (one S.D.) change in own-FTA status has approximately four times the impact on the probability of FTA_{ij} as a typical change in cross-FTA status. These results suggest that the "own-FTA" effect has an economically and statistically larger effect on the explaining FTAs than the "cross-FTA" effect.

Yet, a one S.D. change in $MFTA_{i,t-5}$ (or $MFTA_{j,t-5}$) need not be the same as a one standard deviation change in $ROWFTA_{ij,t-5}$, potentially challenging the conclusion above. Consequently, the last three rows of Table 3a report the marginal response probabilities of a one-unit increase in $MFTA_{i,t-5}$, a one-unit increase in $MFTA_{j,t-5}$ and a two-unit increase for $ROWFTA_{ij,t-5}$. Note that a one-unit—or one-*FTA*—increase in $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) increases the probability of $FTA_{ij,t}$ by 0.44 (0.54) percentage point, and the effect is economically and statistically significant. However, a two-unit increase in $ROW_{ij,t-5}$ increases the probability of $FTA_{ij,t}$ by only 0.01 percentage point, which is neither economically nor statistically significant.[26] Consequently, our conclusion above that the own-FTA effect has an economically and statistically larger effect on explaining FTAs than the cross-FTA is supported strongly.

Table 3a. Marginal Response Probabilities for Natural Trading Partners

Natural trading partners

Existence		95% Confidence interval (C.I.)					*Difference*
		0.1031		0.0995	0.1068		
	$P(FTA\_t = 1 \vert$ CONT = 1)= $\Pr(FTA = 1 \vert X - \sigma)$	95% C.I.		$\Pr(FTA = 1 \vert X + \sigma)$	95% C.I.		$\Pr(FTA = 1 \vert X + \sigma) - \Pr(FTA = 1 \vert X)$
DIFGDP	0.1494	0.1437	0.1551	0.0706	0.0678	0.0734	–0.0325
SUMGDP	0.0716	0.0688	0.0744	0.1474	0.142	0.1529	0.0443
DIST	0.1709	0.1648	0.1769	0.0612	0.0589	0.0636	–0.0419
MDIST	0.0982	0.0943	0.1021	0.1083	0.1041	0.1124	0.0051
MCONT	0.1078	0.1036	0.112	0.0986	0.0948	0.1025	–0.0045
$MFTA_i$	0.0812	0.0782	0.0841	0.1306	0.1259	0.1353	0.0275
$MFTA_j$	0.0766	0.0738	0.0794	0.1381	0.1332	0.143	0.035
ROW FTA	0.0966	0.0929	0.1002	0.1101	0.106	0.1142	0.007
$MFTA_i^*$	0.0989	0.0954	0.1024	0.1075	0.1037	0.1113	0.0044
$MFTA_j^*$	0.098	0.0945	0.1014	0.1085	0.1047	0.1124	0.0054
*ROW FTA**	0.1031	0.0994	0.1067	0.1032	0.0995	0.1068	0.0001

Notes: $\sigma = 1$ S.D. for (first nine variables); σ = count of 1 for $MFTA_i^*$ and $MFTA_j^*$, σ = count of 2 for *ROW FTA**.

Table 3b. Marginal Response Probabilities for Unnatural Trading Partners

Unnatural trading partners

Existence				*95% Confidence interval (C.I.)*			*Difference*
		0.0066		0.0061	0.0071		
	$P(FTA\_t = 1 \mid CONT = 0) = Pr(FTA = 1 \mid X - \sigma)$	95% C.I.		$Pr(FTA = 1 \mid X + \sigma)$	95% C.I.		$Pr(FTA = 1 \mid X + \sigma) - Pr(FTA = 1 \mid X)$
DIFGDP	0.0133	0.0122	0.0143	0.0032	0.0029	0.0034	–0.0034
SUMGDP	0.0032	0.003	0.0035	0.013	0.012	0.014	0.0064
DIST	0.0171	0.0158	0.0183	0.0024	0.0022	0.0026	–0.0042
MDIST	0.006	0.0055	0.0065	0.0072	0.0066	0.0078	0.0006
MCONT	0.0072	0.0066	0.0078	0.006	0.0055	0.0065	–0.0005
$MFTA_i$	0.0041	0.0038	0.0045	0.0103	0.0095	0.0111	0.0037
$MFTA_j$	0.0037	0.0034	0.004	0.0115	0.0106	0.0123	0.0049
ROW FTA	0.0058	0.0053	0.0063	0.0075	0.0069	0.008	0.0009
$MFTA_i^*$	0.0061	0.0056	0.0065	0.0071	0.0066	0.0077	0.0005
$MFTA_j^*$	0.006	0.0055	0.0064	0.0073	0.0067	0.0078	0.0007
*ROW FTA**	0.0066	0.0061	0.0071	0.0066	0.0061	0.0071	0.0000

Notes: $\sigma = 1$ S.D. for (first nine variables); σ = count of 1 for $MFTA_i^*$ and $MFTA_j^*$; σ = count of 2 for *ROW FTA**.

In fact, the increase in the probability of FTA_{ij} of either i or j having one more FTA with another country k (approximately 0.50 percentage point) is 50 times that of the increase in the same probability of one more FTA among another pair kl (approximately 0.01 percentage point). This relative impact is consistent with the relative welfare gains to i or j of 35–60 times of the own-FTA and cross-FTA effects suggested by our numerical theoretical comparative statics.[27]

Table 3b reports the marginal response probabilities for unnatural trading partners (i.e. pairs on different continents). The probability of an FTA among unnatural trading partners at the mean level of all (other) right-hand side variables is 0.0066, with a 95% confidence interval of 0.0061 to 0.0071. The sixth (seventh) line of Table 3b indicates that a one S.D. increase in $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) increases the probability of $FTA_{ij,t}$ to 0.0103 (0.0115). Each of these probability changes is statistically significant at the 95% level, but not significantly different from one another. By contrast, a one S.D. increase in $ROWFTA_{ij,t-5}$ increases the probability of $FTA_{ij,t}$ to 0.0075, which is not a statistically significant change. The difference in the marginal response probabilities for $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) and $ROWFTA_{ij,t-5}$ is economically and statistically significant. Moreover, the difference in such probabilities is as expected; a one S.D. change in $MFTA_{i,t-5}$ ($MFTA_{j,t-5}$) has a quantitatively larger impact on the likelihood of $FTA_{ij,t}$ than does a one S.D. change in $ROWFTA_{ij,t-5}$. For brevity, we do not review the other marginal response probabilities. However, all such probabilities change in the expected directions and all such changes are statistically significant except for $MCONT_{ij}$. These results also suggest that the "own-FTA" effect has an economically and statistically larger effect on explaining FTAs than the "cross-FTA" effect.

Finally, we note that the change in the probability of FTA_{ijt} owing to an increase in $MFTA_{i,t-5}$ (or $MFTA_{j,t-5}$ or $ROWFTA_{ij,t-5}$) is much higher for natural trading partners than for unnatural trading partners. For instance, for $MFTA_{i,t-5}$ a one unit increase in this variable (that is, one more bilateral FTA for i) causes a 0.44 percentage point (0.05 percentage point) increase in the likelihood of FTA_{ij} within the next 5 years if i and j are on the same (a different) continent. This is one approach for circumventing distance-weighted measures of $MFTA$ and $ROWFTA$.

5. Predicting FTAs

An alternative measure of goodness-of-fit for logit and probit models is the "percentage correctly predicted." However, Wooldridge (2000) points out that this percentage may be misleading. For instance, in BB, the authors had a sample of 1431 country pairs for the year 1996 with 286 actual FTAs (true positives, or TPs). Hence, 20% of the observations were FTAs. The "unconditional" probability of an FTA was 20% and the unconditional probability of No-FTA was 80% (1145/1431). Consequently, even if the model had no explanatory power and failed to predict correctly even one FTA, the percentage of No-FTAs correctly predicted is almost 80%. This large percentage misrepresents the zero predictive power of the model for predicting true positives.

Wooldridge (2000) recommends examining separately the percentage correctly predicted for each of the two outcomes. That is, the percentage of "true positives" (TPs) in "all positives" (APs), or TPs/APs = TPs/(TPs + FPs) where FP denotes "false positives," is important, but so is the percentage of "true negatives" (TNs) in "all negatives" (ANs), or TNs/ANs = TNs/(TNs + FNs) where FN denotes "false negative." BB conducted this statistical summary for their cross-section analysis of year 1996 data and found that their model predicted correctly 243 of 286 FTAs, or 84.97%. They also predicted 1114 of the 1145 pairs without FTAs correctly, or 97.29%.

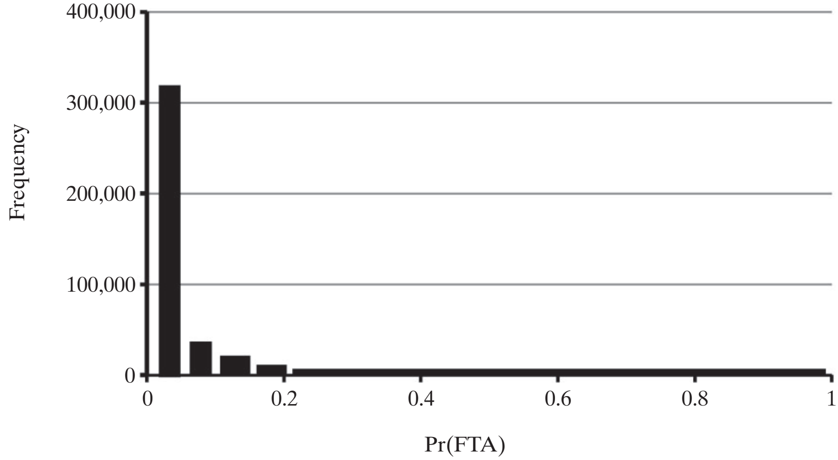

Figure 3a. Frequency of Predicted Probabilities (Existence, Logit, Simple Average Weight)

However, a critical issue in classification is the choice of the "cutoff" on the probability continuum. BB, EL and CJ followed McFadden (1975, 1976) in using a probability cutoff (p^C) of 0.5 to determine if an FTA was predicted or not. Letting p_{ij} denote the predicted probability from the probit regression in BB, if $p_{ij} > 0.5$ and the country-pair *ij* had an FTA, this would be a true positive. If $p_{ij} \leq 0.5$ and *ij* did not have an FTA, this would be a true negative.

In this part, we examine some summary statistics associated with alternative cutoff probabilities. We examine four alternative methods for assessing the overall predictive power of our main logit models for existence of and transition to FTAs, which are Specifications 3 and 4, respectively; both specifications exclude any fixed effects. The first concerns establishing a cutoff probability based upon maximizing the overall predictive power; this is determined by a "specificity–sensitivity" analysis, described shortly. The second and third concern establishing cutoff probabilities consistent with having a TN rate no lower than that in BB (97%) or EL (99%), respectively.[28] The fourth uses the arbitrary cutoff of 0.5, but it turns out that this cutoff is consistent with a true negative rate of 99% (as in the third approach).

While BB, EL and CJ used a p^C of 0.5, we believe this cutoff is not a very relevant one. The reason lies in the fact that—as noted earlier—bilateral FTA events in our panel of over 350,000 observations are rare events. First, the number of observations when an FTA exists between a country-pair in a given year is 10,478; this is only 3% of all observations. Second, the number of observations when a country-pair forms (or transitions to) an FTA is 3811; this is only 1% of all observations. Figure 3a provides a plot of the frequency of the predicted probability of an FTA (p_{ijt}) using Specification 3; this confirms visually that FTAs are rare events and that a $p_{ijt} > 0.5$ would be an extremely rare event. Consequently, we ignore this cutoff for now, although for completeness we will report the TP and TN rates for $p^C = 0.5$ later.

Cohen et al. (2003) suggests using *a priori* information about the proportion of FTA events and No-FTA events in our population. Consider first the case of FTA existences. The proportion of FTAs (No-FTAs) in our panel—which is virtually the entire population of country-pairs since 1960—is 3% (97%). Hence, the unconditional probability of an FTA existing between any country-pair in a given year is 3%. This sug-

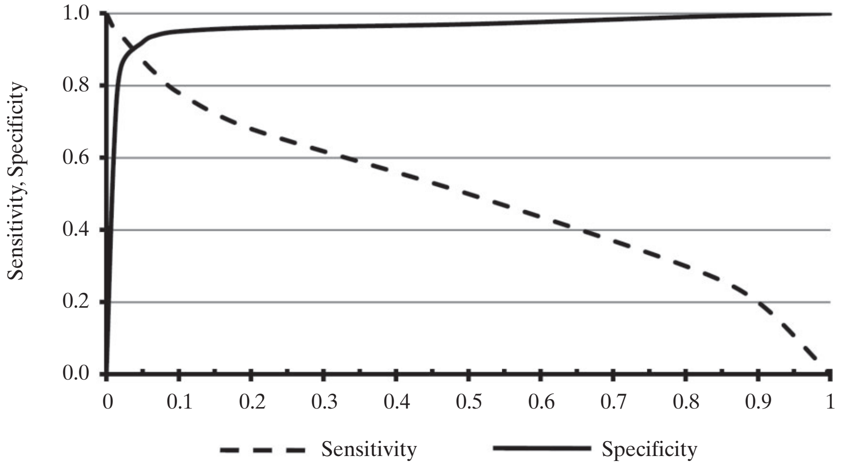

Figure 3b. Sensitivity, Specificity (Existence, Logit, Simple Average Weight)

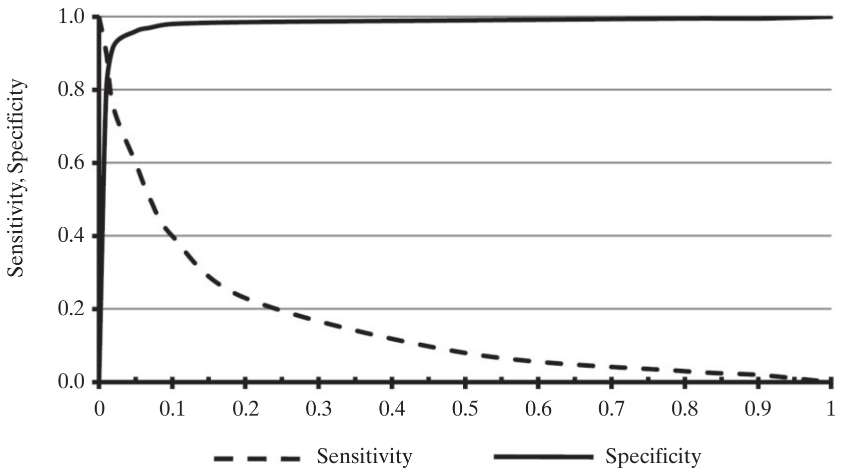

Figure 3c. Sensitivity, Specificity (Transitions, Logit, Simple Average Weight)

gests a more appropriate cutoff probability is 0.03; BJ followed this approach also. In fact, it turns out that the TP and TN rates are maximized at this cutoff, as we now show.

Naturally, one wants to maximize both the rates of true positives (TPs) and true negatives (TNs). However, there is a trade-off. Figure 3b graphs the TP and TN rates for Specification 3 (logit, FTA existence) against the entire range of possible cutoff probabilities; Figure 3c graphs the TP and TN rates against the same range of p^C for Specification 4 (logit, FTA transition). One can see from Figure 3b that at p^C of near 0, one maximizes the likelihood of predicting an FTA when one exists (sensitivity); however, the TN rate (specificity) is virtually zero, which is a severe problem since the vast bulk of observations is zero. To increase the TN rate, a higher p^C is needed. For our first approach, it turns out that at a p^C of 0.03 (specifically, 0.0328) we maximize both the TP and TN rates at 91%; this is the result from a sensitivity–specificity analysis. Thus, at a cutoff probability consistent with the unconditional probability of an

FTA existing (0.03), the model predicts correctly 91% of the cases when an FTA exists within 5 years of the agreement forming and 91% of the cases when No-FTA is correct. We have also conducted this analysis for predicting formations of FTA ($TFTA_{ijt}$ equals 1 in the year an FTA goes into force, and 0 otherwise). In this case, the model predicts correctly 89% of the true positives within 5 years of the formation and 89% of the TNs, as shown in Figure 3c.

Our second and third approaches consider two other possible cutoff probabilities. In the first approach, we obtain a success rate for predicting FTAs when FTAs exist of 91% (the TP rate). While the TN rate may seem high at 91%, we still have a false positive rate of 9%, which implies in our sample of over 350,000 observations that we incorrectly predict FTA when No-FTA exists in 9% of the cases. However, BB had a higher TN rate of 97% (owing to its $p^C = 0.5$) and EL had a TN rate of 99% (also using a $p^C = 0.5$), implying much stricter false positive rates of 3% and 1%, respectively. As Figure 3b suggests, one can raise the p^C to ensure a higher TN rate, to be consistent with these studies, which will of course lower the TP rate. We considered two alternative values of p^C. First, we considered $p^C = 0.114$, which ensured a TN rate of 97% as in BB. The associated TP rate was 75%. The latter value is only 10% less than the 85% TP rate in BB for only a cross-section of bilateral FTAs among 53 country-pairs. Our TP rate of 75% is remarkably high considering we are predicting the existence of an FTA between a country-pair within only 5 years of its formation. For a TN rate of 99% (implying $p^C = 0.307$), the TP rate for existence of an FTA between a country-pair within 5 years of its formation falls to 48%. We also considered the TP rates for predicting the actual year of formation (date of entry) of an FTA between a country-pair within 5 years of its actual formation. At a TN rate of 97%, the TP rate is 56%. At a TN rate of 99%, the TP rate is 26%. Our fourth approach simply uses the cutoff probability of 0.5.

Tables 4a and b summarize the information above and additionally provide information about the TP and TN rates by individual year as well as with and without the $MFTA_{i,t-5}$, $MFTA_{j,t-5}$ and $ROWFTA_{ij,t-5}$ terms. For economy, we provide the predictions at 5-year intervals as well as over all the years, where the logit specification in Table 4a includes the *MFTA* and *ROWFTA* terms and the specification in Table 4b excludes these terms. First, in the second and third columns, we use the cutoff that maximizes overall success rate using a sensitivity–specificity analysis, that is, maximizing both the TP and TN rates. For FTA existences including *MFTA* and *ROWFTA*, this is 91.04% in Table 4a. In Table 4b, we can see from columns (2) and (3) that the percentage correctly predicted without *MFTA* and *ROWFTA* is 88.35%. However, returning to Table 4a, this 91% still leaves 9% of the observations false negatives. The second approach considered a TN rate no lower than that in BB, 97%. With a higher TN rate, the fourth and fifth columns report a lower TP rate of 75.50% (65.34%) in the specification with (without) the *MFTA* and *ROWFTA* terms. In our third approach with a TN rate of 99% as in EL, the TP rate falls to 47.67% (32.37%) in the specification with (without) the *MFTA* and *ROWFTA* terms. Fourth, the eighth and ninth columns provide the TP rates using a cutoff of $p^C = 0.5$. The TP rate is 47.07%, which is similar to that using the 99% TN rate in the third approach. However, by contrast with the logit specification omitting the *MFTA* and *ROWFTA* terms, the predictive power of this logit is better; in Table 4b we predict only 28.31% of the FTA cells correctly using a cutoff of 0.5.[29] One more interesting result is worth noting from a comparison of Tables 4a and b. In the case of Table 4a, the presence of the *MFTA* and *ROWFTA* terms in the specification causes the percentage correctly predicted to increase as time progresses; however, in the case of Table 4b, the percentage correctly predicted falls over time. Hence,

Table 4a. Predictions of FTAs

	Correctly predict FTA existence in year t (FTA_t = 1) using MFTA and ROWFTA variables								
	Cutoff from Sens, Sp		*TrueNeg.Rate = 97%*		*TrueNeg.Rate = 99%*		*Prob. Cutoff at 50%*		
Year t	*%t Correct*	*Correctly predicted*	*% Correct*	*Correctly predicted*	*% Correct*	*Correctly predicted*	*% Correct*	*Correctly predicted*	*Observed*
1960	83.87	26	38.71	12	30.96	1	3.23	1	31
1965	75.44	43	56.14	32	56.98	6	10.53	6	57
1970	82.93	34	65.85	27	19.51	8	17.07	7	41
1975	86.54	90	62.50	65	22.12	23	21.15	22	104
1980	88.46	92	84.62	88	56.73	59	56.73	59	104
1985	87.22	116	75.94	101	46.62	62	45.86	61	133
1990	94.30	149	81.65	129	57.59	91	56.96	90	158
1995	93.96	280	76.85	229	45.97	137	44.30	132	298
2000	91.79	570	76.17	473	50.40	313	50.08	311	621
2005	91.94	867	76.78	724	51.64	487	51.22	483	943
Total	91.04	2267	75.50	1880	47.67	1187	47.07	1172	2490
Cutoff		0.039		0.1344		0.49		0.5	

Notes: The following predictions are from Logit, MFTA count regression, with years in multiples of 5 only. Independent variables are in t – 5.

Table 4b. Predictions of FTAs

	Correctly predict FTA existence in year t (FTA_t = 1) without MFTA and ROW FTA variables								
	Cutoff from Sens, Sp		*TrueNeg.Rate = 97%*		*TrueNeg.Rate = 99%*		*Prob. Cutoff at 50%*		
Year t	*% Correct*	*Correctly predicted*	*% Correct*	*Correctly predicted*	*% Correct*	*Correctly predicted*	*% Correct*	*Correctly predicted*	*Observed*
1960	100.00	31	90.32	28	51.61	16	41.94	13	31
1965	89.47	51	61.40	35	29.82	17	26.32	15	57
1970	97.56	40	73.17	30	48.78	20	41.46	17	41
1975	95.19	99	81.73	85	45.19	47	38.46	40	104
1980	98.08	102	82.69	86	47.12	49	43.27	45	104
1985	98.50	131	81.95	109	43.61	58	38.35	51	133
1990	98.10	155	83.54	132	44.94	71	40.51	64	158
1995	98.32	293	79.19	236	40.94	122	34.23	102	298
2000	88.24	548	63.93	397	29.95	186	25.76	160	621
2005	79.53	750	51.86	489	23.33	220	21.00	198	943
Total	88.35	2200	65.34	1627	32.37	806	28.31	705	2490
Cutoff		0.027		0.1368		0.45		0.5	

Notes: The above predictions are from Logit, (*No-MFTA*) count regression, with years in multiples of 5 only. Independent variables are in $t - 5$.

accounting for endogenous bilateralism in the logit specification contributes to a relatively more successful true positive rate over time.

6. Conclusions

One of the most notable international economic events since 1990 has been the proliferation of bilateral FTAs, argued by some to be attributable to governments having pursued a policy of "competitive liberalization." Guided by Baier and Bergstrand (2004), we have explored the relative importance for the welfare gains of an FTA between a country-pair *ij* (FTA_{ij}) of two "sources" of interdependence, own-FTA interdependence (the effect of an existing FTA of *i* or *j* with a third country *k*) and cross-FTA interdependence (the effect of an existing FTA between a third-country-pair *kl*). Theoretical considerations suggest that the own-FTA effect of one more FTA has an impact on the gains from FTA_{ij} greater than that of one more FTA among a third-country-pair. Guided by these considerations, we specified a simple logit (and probit) model to estimate the influence on the likelihood of a bilateral FTA between *i* and *j* of indexes for each country of "multilateral FTAs" and "*ROW* FTAs"—in the spirit of Anderson and van Wincoop's (2003) "multilateral resistance" terms, as linearized in Baier and Bergstrand (2009). We found that the marginal response probabilities of these indexes of "own-FTA" and "cross-FTA" effects were both statistically and economically significant, and the response probabilities suggested that one more FTA of *i* or *j* with a third country *k* (own-FTA effect) had an impact on the probability of FTA_{ij} of approximately 50 times that of one more FTA among a third-country-pair *kl* (cross-FTA effect). Moreover, using a "sensitivity–specificity" analysis, we determined the optimum cutoff probability for predicting FTAs and the results indicated that we could predict correctly an FTA ("No-FTA") when one existed (none existed) 91% of the time. The results provide economically and statistically significant evidence that own-FTA effects tend to dominate cross-FTA effects as sources of "interdependence."

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Notes

1. According to Heydon and Woolcock (2009, pp. 10–11), bilateral preferential trade agreements "account for 80% of all PTAs notified and in force; 94% of those signed and/or under negotiation; and 100% of those at the proposal stage" with the vast bulk of preferential trade agreements being FTAs. Heydon and Woolcock note that "among projected agreements" 92% are planned as FTAs, 7% as partial scope agreements and only 1% as customs unions, with customs unions differing from FTAs owing to the former having a common external tariff with nonmembers. For brevity, we refer here to FTAs and customs unions as "FTAs," as most agreements formed in the past 50 years have been FTAs. Our analysis will omit partial scope agreements.

2. The reason for this exclusion is that BB examined only a cross-section for a particular year. Evaluating that issue econometrically would have led to endogeneity issues beyond that paper's scope, necessitating the computationally demanding cross-sectional spatial econometric techniques applied in Egger and Larch (2008).

3. In their panel data estimation, the spatial lag of third-country-pairs' FTAs were time-lagged 5 years to avoid endogeneity. Also, they combined own-FTA and cross-FTA influences by allowing $kl = il$ and $kl = kj$. Moreover, to avoid possible correlation of time-varying variables with the time-invariant component of the error terms, EL employed Chamberlain (1980) time-meaned "fixed effects." We discuss implications of this later.

4. The theoretical frameworks in Baldwin and Jaimovich (2012) and Chen and Joshi (2010), to be discussed shortly, are limited to three countries and so can only address the effects of an FTA between countries k and i or k and j on the welfare gains of an FTA between i and j. Egger and Larch (2008) did not provide a formal theoretical model. The underlying theoretical framework in BB has six countries.

5. The term "tariff complementarity" was introduced in Bagwell and Staiger (1998). Tariff complementarity (static version) refers to the net welfare gain to a member of an existing FTA from lowering external tariffs on a nonmember. Neither EL nor BJ addressed tariff complementarity, focusing instead upon trade diversion of nonmembers. This will be discussed more later. Some numerical comparative statics presented in an Online Appendix based upon the BB model suggest that the own-FTA welfare effect of one more FTA exceeds the cross-FTA welfare effect by a magnitude of 35–60 times. For the Online Appendix see Supporting Information.
6. The FTA indexes are akin to (inverse) multilateral and ROW resistances in the structural gravity model in Anderson and van Wincoop (2003), as linearized using a Taylor-series expansion in Baier and Bergstrand (2009).
7. We use "count" variables. CJ does not capture the full influence of own-FTA effects as we do using count variables; instead, CJ used indicator (dummy) variables.
8. The models in EL and BJ employed fixed effects.
9. In the case of one industry, there are only two parameters in the BB model, the elasticity of substitution in consumption (σ) and the level of fixed costs in production (ϕ). We assume that the countries have identical absolute (and hence relative) factor endowments, the elasticity of substitution between varieties (σ) is 4, the level of fixed costs is set at 1 (without loss of generality), and tariff rates are initially 30% for all countries ($\tau = 0.3$), as in BB. The sensitivity of the numerical comparative statics to values of σ and τ are discussed in the Online Appendix (see Supporting Information); the results are insensitive qualitatively to the level of fixed costs, ϕ.
10. In the Online Appendix, we show explicitly in the context of the BB model that the welfare gain for $1A$ from the formation of $FTA_{1A,1B}$—conditioned upon the existence of $FTA_{2A,2B}$—is unambiguously higher than the welfare gain for $1A$ from the formation of $FTA_{1A,1B}$ in the absence of $FTA_{2A,2B}$ for $\sigma = 4$ and initial $\tau = 0.3$.
11. We refer to the incentive for $1A$ to form an FTA with $1B$, conditioned on $FTA_{1A,2A}$, as tariff complementarity; this is a "selective" form of tariff complementarity. Typically, tariff complementarity refers to MFN tariffs (cf. Bagwell and Staiger (1998)). Here, tariff complementarity does not address MFN external tariffs but rather a specific form of tariff liberalization, namely, another FTA. However, this is similar, but not identical, to using Nash equilibrium tariffs, the typical setting for discussing (static) tariff complementarity.
12. In our Online Appendix, we show explicitly in the context of the BB model that the welfare gains from $FTA_{1A,1B}$ conditioned upon $FTA_{1A,2A}$ are approximately 35–60 times the welfare gains from $FTA_{1A,1B}$ conditioned upon $FTA_{2A,2B}$ for $\sigma = 4$ and initial $\tau = 0.3$. We also show why $2A$ retains its incentive to stay in the agreement with $1A$.
13. We will also consider probit estimates for robustness. However, as will be discussed later, logit is less restrictive (and problematic) than probit for introducing fixed effects in the robustness analysis.
14. In the Online Appendix we show these measures are in accordance with our numerical comparative statics. In Appendix Figures 5a–b, we introduce one FTA between $2A$ and $2B$ to see its impact on the net welfare gains of an FTA between $1A$ and $1B$. In Appendix Figures 6a–c, we introduce one FTA between $1A$ and $2A$ to see its impact on the net welfare gains of an FTA between $1A$ and $1B$.
15. Theory cannot suggest the specific choice of the 5-year time lag. However, two considerations influenced the decision. First, the decision by a country pair to form an FTA takes considerable time, and often exceeds 5 years. Second, in many cases, interdependence may cause FTA formations to accelerate and consequently 5 years may be too short a period. We chose 5 years to balance both sets of considerations.
16. Of course, GDPs increase over time and consequently not scaling GDPs of countries by world GDP may influence the results. Consequently, we also considered weights $\theta_{k,t-5} = Y_{k,t-5}/Y^{W}_{t,t-5}$, where Y^{W}_{t-5} is world GDP. The results are robust to this alternative measure. The alternative measure only influences the absolute magnitudes of the coefficient estimates, but has no bearing on the marginal response probabilities.

17. The data base is available at www.nd.edu/ jbergstr. Documentation for its construction is provided at the website. Every positive cell entry is hyper-linked to a PDF of its original treaty (98% of cells) or a secondary source (2% of cells); not all cells are potential observations as over the period some countries formed and others dissolved, e.g. Czechoslovakia. We will use only 146 of these countries, as explained shortly.
18. In the cases of these variables, we use averages rather than "count" variables, for convenience; this has no material consequence for the results.
19. We also used $ln[(Y_{it}+Y_{jt})/Y_t^W]$ for economic size, since world GDP changes over time; the results are robust to this alternative measure. Also, as in BB and EL, we are using real GDPs as a proxy for absolute factor endowments. Consequently, the terms-of-trade (real income) effects from, say, a natural FTA relative to an unnatural FTA are captured by $DIST_{ij}$ and $CONT_{ij}$.
20. There is likely a high degree of persistence in FTAs; the effect of this persistence is mitigated by predicting instead "formations."
21. A similar fall in overall explanatory power for the same adjustment was found in EL.
22. We also ran the probit on the FTA transitions dummy ($TFTA_{ij,t}$) and the coefficient estimates are qualitatively identical to those using the corresponding logit specification, but not reported for brevity. As demonstrated in Cameron and Trivedi (2005, p. 473), probit estimates should be approximately 55% of the corresponding logit estimates ($0.55=\sqrt{3}/\pi$).
23. Since Specification 6 is the only one in Table 2 to use the GDP-weighted versions of $MFTA_{i,t-5}$, $MFTA_{j,t-5}$, $ROWFTA_{ij,t-5}$, $MDIST_{ij,t-5}$ and $MCONT_{ij,t-5}$, we do not change the names of the variables named in column (1) to keep Table 2's size manageable.
24. This conclusion also holds for the FTA transitions LHS variable ($TFTA_{ij,t}$). Using GDP weights causes $MDIST_{ij}$ and $MCONT_{ij}$ to become time-varying. Also, see note 23 regarding the effects of the alternative weights on empirical results. Coefficient estimates can change but marginal response probabilities are unaffected. Hence, the larger coefficient estimate for *ROWFTAY* in Specification 6 does not alter significantly its marginal response probability.
25. Note that each country pair enters the data set only *once*, unlike gross trade flows in gravity equations. Thus, the coefficients on $MFTA_{i,t-5}$ and $MFTA_{j,t-5}$ need not be exactly equal; if every pair entered twice, they would be exactly equal.
26. In the context of Hypothesis 1, conditioning on $FTA_{1A,2A}$ is equivalent to a one-unit increase whereas—in the context of Hypothesis 2—conditioning on $FTA_{2A,2B}$ is equivalent to a two-unit increase.
27. These numerical comparative static estimates are provided in the Online Appendix. Note, however, that our marginal response probabilities—calculated using either a one-standard-deviation or a one-unit change—are employed to evaluate empirically (as closely as feasible) our theoretical Hypotheses 1 and 2. However, one could argue that the number of third-country-pair FTAs that a typical country-pair faces combined with the $ROWFTA_{ij,t-5}$ marginal response probability should be compared with the number of own FTAs combined with the $MFTA_{i,t-5}$ marginal response probability. We leave this to future research; our theoretical model only provides predictions concerning unit increments in other FTAs.
28. However, the predicted probabilities in EL are based upon specifications using time-meaned fixed effects.
29. The near doubling in predictive TP rates is much larger than the 5 percentage point improvement in Egger and Larch (2008) from introducing their spatial lag.

Supporting Information

Additional Supporting Information may be found in the online version of this article at the publisher's web-site:

SI. Online Appendix.