Journal of International Economics 71 (2007) 72 – 95
Journal of INTERNATIONAL ECONOMICS
www.elsevier.com/locate/econbase

Do free trade agreements actually increase members' international trade?

Scott L. Baier [a,b], Jeffrey H. Bergstrand [c,d,*]

[a] John E. Walker Department of Economics, Clemson University, Clemson, SC 29634, USA
[b] Federal Reserve Bank of Atlanta, Atlanta, GA 30309, USA
[c] Department of Finance, Mendoza College of Business, and Kellogg Institute for International Studies, University of Notre Dame, Notre Dame, IN 46556, USA
[d] CESifo, Germany

Received 24 April 2003; received in revised form 15 December 2005; accepted 15 February 2006

Abstract

For over 40 years, the gravity equation has been a workhorse for cross-country empirical analyses of international trade flows and — in particular — the effects of free trade agreements (FTAs) on trade flows. However, the gravity equation is subject to the same econometric critique as earlier cross-industry studies of U.S. tariff and nontariff barriers and U.S. multilateral imports: trade policy is *not* an exogenous variable. We address econometrically the endogeneity of FTAs. Although instrumental-variable and control-function approaches do not adjust for endogeneity well, a panel approach does. Accounting econometrically for the FTA variable's endogeneity yields striking empirical results: the effect of FTAs on trade flows is quintupled. We find that, on average, an FTA approximately *doubles* two members' bilateral trade after 10 years.

Keywords: Free trade agreements; International trade flows; Gravity equation

JEL classification: F1

1. Introduction

> The issue of exogeneity may also be an important problem when dummy variables are used (in a gravity equation) to estimate the effects of free trade areas (Lawrence, 1998, p. 59).

* Corresponding author. Department of Finance, Mendoza College of Business, and Kellogg Institute for International Studies, University of Notre Dame, Notre Dame, IN 46556, USA. Tel.: +1 574 631 6761; fax: +1 574 631 5255.
E-mail address: Bergstrand.1@nd.edu (J.H. Bergstrand).

0022-1996/$ - see front matter
doi:10.1016/j.jinteco.2006.02.005

One might expect — having witnessed a virtual explosion in the number of free trade agreements (FTAs) among nations over the past decade and a half — that the answer to the question posed in this paper's title is unequivocal: *yes*! Surprisingly, international trade economists can actually claim little firm empirical support for reliable quantitative estimates of the average effect of an FTA on bilateral trade (all else constant).

Over the past 40 years, the "gravity equation" has emerged as the empirical workhorse in international trade to study the ex post effects of FTAs and customs unions on bilateral merchandise trade flows.[1] The gravity equation is typically used to explain cross-sectional variation in country pairs' trade flows in terms of the countries' incomes, bilateral distance, and dummy variables for common languages, for common land borders, and for the presence or absence of an FTA. Nobel laureate Jan Tinbergen (1962) was the first to publish an econometric study using the gravity equation for international trade flows, which included evaluating the effect of FTA dummy variables on trade. His results suggested economically insignificant "average treatment effects" of FTAs on trade flows. Tinbergen found that membership in the British Commonwealth (Benelux FTA) was associated with only 5 (4) percent higher trade flows. Since then, results have been mixed, at best. For example, Aitken (1973), Abrams (1980), and Brada and Mendez (1985) found the European Community (EC) to have an economically and statistically significant effect on trade flows among members, whereas Bergstrand (1985) and Frankel, Stein and Wei (1995) found insignificant effects. Frankel (1997) found positive significant effects from Mersosur, insignificant effects from the Andean Pact, and significant *negative* effects from membership in the EC in certain years. He noted:

> If the data from four years — 1970, 1980, 1990, 1992 — are pooled together, the estimated coefficient on the European Community is a smaller 0.15, implying a 16 percent effect" (p. 83).

Frankel (1997) concluded that several readers "have found surprising our result that intra-European trade can be mostly explained by various natural factors, with little role for the EC until the 1980s…." (p. 88). Other studies in international trade have had similar seemingly implausible results.[2]

The fragility of estimated FTA treatment effects is addressed directly in Ghosh and Yamarik (2004). These authors use extreme-bounds analysis to test the robustness of FTA dummy coefficient estimates. They find empirical evidence using cross-section data that the estimated average treatment effects of most FTAs are "fragile," supporting our claims. Thus, there still are no reliable ex post estimates of the FTA average treatment effect. This paper is aimed at addressing this puzzle.

All these studies, however, typically assume an *exogenous* right-hand-side (RHS) dummy variable to represent the FTA treatment. In reality, FTA dummies are not exogenous random variables; rather, countries likely select endogenously into FTAs, perhaps for reasons unobservable to the econometrician and possibly correlated with the level of trade.[3] This paper applies developments in the econometric analysis of treatment effects — some well-known and others more recent — to estimate the effects of FTAs on bilateral trade flows using a panel of cross-

[1] Bayoumi and Eichengreen (1997, p. 142) note that the gravity equation has "long been the workhorse for empirical studies of the pattern of trade." This study (purposefully) does not address ex ante analyses of the effects of FTAs on trade flows using computable general equilibrium models.

[2] Frankel (1997) and Oguledo and MacPhee (1994) provide summaries of FTA coefficient estimates across studies. Frankel (1997, pp. 86–90) draws considerable attention to the surprising insignificant effects (especially prior to the 1980s) of the EC and EFTA in his and others studies, such as Bergstrand (1985, 1989) and Boisso and Ferrantino (1997). However, no systematic explanation is provided.

[3] We note that, for about a decade, several researchers have acknowledged potential endogeneity bias, but only that created by GDPs as RHS variables. Several authors have instrumented for GDPs, but (with the exception of the three studies noted shortly) none have instrumented for FTAs.

section time-series data at five-year intervals from 1960 to 2000 for 96 countries. The literature on treatment effects, developed in the context of numerous labor economics studies (cf., Wooldridge, 2002), provides rich tools that have not previously been used for analyzing the effects of bilateral trade policies on international trade flows.

This is not the first paper in empirical international trade to call attention to the potential endogeneity bias in estimating the effect of trade policies on trade volumes. For instance, Trefler (1993) addressed systematically the simultaneous determination of U.S. multilateral imports and U.S. multilateral nontariff barriers in a cross-industry analysis. Trefler found using instrumental variables that, after accounting for the endogeneity of trade policies, the effect of these policies on U.S. imports increased *tenfold*. Lee and Swagel (1997) also showed using instrumental variables that previous estimates of the impact of trade liberalization on imports had been considerably underestimated.

Clearly, the literature on bilateral trade flows and bilateral FTAs using the gravity equation is subject to the same critique that Trefler raised: the presence or absence of an FTA is not exogenous. The issue is important because – if FTAs are endogenous – previous cross-section empirical estimates of the effects of FTAs on trade flows may be biased, and the effects of FTAs on trade may be seriously over- or under-estimated, as the extreme-bounds evidence in Ghosh and Yamarik (2004) suggests. To date, only three papers have attempted to address the potential bias in cross-section gravity models caused by endogenous FTAs, Baier and Bergstrand (2002, 2004b) and Magee (2003). However, all three papers — using instrumental variables with cross section data — provide at best mixed evidence of isolating the effect of FTAs on trade flows.

The empirical results in our paper suggest three important conclusions. First, several plausible reasons exist to suggest that the quantitative (long-run) effects of FTAs on trade flows using the standard cross-section gravity equation are biased; we argue that unobservable heterogeneity most likely biases estimates *downward*. Second, we find that, owing to this bias, traditional estimates of the effect of FTAs on bilateral trade flows have tended to be underestimated by as much as 75–85%. Third, we demonstrate that the most plausible estimates of the average effect of an FTA on a bilateral trade flow are obtained from a theoretically-motivated gravity equation using panel data with bilateral fixed and country-and-time effects or differenced panel data with country-and-time effects. Other methods to identify the impact, such as instrumental variables using cross-section data, are compromised by a lack of suitable instruments. We find that, on average, an FTA approximately *doubles* two members' bilateral trade after 10 years.

Section 2 presents a (traditional) atheoretical cross-section gravity equation and a (modern) theoretically-motivated cross-section gravity equation. Section 3 provides motivation for suspecting endogeneity of FTA dummy variables and suggests the likely direction of bias. Section 4 summarizes some instrumental-variable and control-function econometric studies that have tried to eliminate endogeneity bias using cross-section data. Section 5 addresses panel techniques and discusses the results from applying various fixed and time effects and first differencing to panel data. Section 6 concludes.

2. The gravity equation in international trade

The gravity equation in international trade most commonly estimated using cross-country data is:

$$\mathrm{PX}_{ij} = \beta_0(\mathrm{GDP}_i)^{\beta 1}(\mathrm{GDP}_j)^{\beta 2}(\mathrm{DIST}_{ij})^{\beta 3}\mathrm{e}^{\beta 4(\mathrm{LANG}_{ij})}\mathrm{e}^{\beta 5(\mathrm{ADJ}_{ij})}\mathrm{e}^{\beta 6(\mathrm{FTA}_{ij})}\varepsilon_{ij} \tag{1}$$

where PX_{ij} is the value of the merchandise trade flow from exporter i to importer j, GDP_i (GDP_j) is the level of nominal gross domestic product in country i (j), DIST_{ij} is the distance between the economic centers of countries i and j, LANG_{ij} is a binary variable assuming the value 1 if i and j share a common language and 0 otherwise, ADJ_{ij} is a binary variable assuming the value 1 if i and j share a common land border and 0 otherwise, FTA_{ij} is a binary variable assuming the value 1 if i and j have a free trade agreement and 0 otherwise, e is the natural logarithm base, and ϵ_{ij} is assumed to be a log-normally distributed error term.[4]

The earliest applications of the gravity equation to international trade flows were not grounded in formal theoretical foundations, cf., Tinbergen (1962), Linnemann (1966), Aitken (1973) and Sapir (1981); these earlier studies appealed either to informal economic foundations or to a physical science analogy. Since 1979, however, formal theoretical economic foundations for a gravity equation similar to Eq. (1) have surfaced, cf., Anderson (1979), Bergstrand (1985), Deardorff (1998), Baier and Bergstrand (2001), Eaton and Kortum (2002), and Anderson and van Wincoop (2003). A notable feature common to *all* these models is an explicit role for *prices*; in all six papers, price levels or some form of multilateral price indexes surface theoretically.[5] Most recently, Anderson and van Wincoop (2003) illustrated the omitted variables bias introduced by ignoring prices in the cross-section gravity equation. Their framework suggests theoretically the gravity model should be estimated as:

$$\begin{aligned}\ln[\text{PX}_{ij}/(\text{GDP}_i\text{GDP}_j)] = \beta_0 + \beta_3(\ln\text{DIST}_{ij}) + \beta_4(\text{ADJ}_{ij}) + \beta_5(\text{LANG}_{ij}) \\ + \beta_6(\text{FTA}_{ij}) - \ln P_i^{1-\sigma} - \ln P_j^{1-\sigma} + \varepsilon_{ij}\end{aligned} \tag{2}$$

subject to N equilibrium conditions:

$$P_1^{1-\sigma} = \sum_{i=1}^{N} P_i^{\sigma-1}(\text{GDP}_i/\text{GDP}^{\text{W}})e^{\beta 3(\ln\text{DIST}_{i1})+\beta 4(\text{LANG}_{i1})+\beta 5(\text{ADJ}_{i1})+\beta 6(\text{FTA}_{i1})} \tag{3.1}$$

•
•
•

$$P_N^{1-\sigma} = \sum_{i=1}^{N} P_i^{\sigma-1}(\text{GDP}_N/\text{GDP}^{\text{W}})e^{\beta 3(\ln\text{DIST}_{iN})+\beta 4(\text{LANG}_{iN})+\beta 5(\text{ADJ}_{iN})+\beta 6(\text{FTA}_{iN})} \tag{3.N}$$

to generate unbiased estimates of β_0, β_3, β_4, β_5 and β_6. GDP^{W} denotes world GDP, constant across countries. $P_i^{1-\sigma}$ and $P_j^{1-\sigma}$ are denoted "multilateral (price) resistance terms."[6] Anderson and van Wincoop then estimate this system using a custom nonlinear least squares program, treating all $P_i^{1-\sigma}$ variables ($i=1,\ldots, N$ countries) as endogenous. However, Anderson and van Wincoop (2003) and Feenstra (2004, Ch. 5) both suggest that an alternative — and computationally easier — method for accounting for multilateral price terms $P_i^{1-\sigma}$ and $P_j^{1-\sigma}$ in cross section — that will also generate

[4] The standard gravity equation sometimes includes the exporter and importer populations or per capita GDPs. Theoretical foundations for this alternative specification still require refinement, which is beyond the scope of this particular paper but is an issue addressed in Bergstrand (1989, 1990) and Baier and Bergstrand (2002).

[5] A seminal contribution to the theoretical foundations of the gravity equation, but in the absence of trade costs (and consequently price terms), is Helpman and Krugman (1985, Ch. 8).

[6] Assuming distance, language, adjacency and FTA are *symmetric* for *ij* and *ji*, the equilibrium multilateral resistance terms can be solved for according to Eqs. (3.1) through (3.N).

Table 1
Typical cross-section gravity equation coefficient estimates

Variable	(1) 1960	(2) 1970	(3) 1980	(4) 1990	(5) 2000
ln GDP_i	0.76 (45.79)	0.88 (57.55)	1.01 (69.37)	1.08 (85.13)	1.18 (104.13)
ln GDP_j	0.76 (48.66)	0.92 (63.95)	1.00 (72.69)	0.97 (78.08)	0.98 (87.39)
ln $DIST_{ij}$	−0.64 (−16.23)	−0.85 (−21.10)	−1.06 (−28.15)	−1.07 (−28.82)	−1.17 (−32.57)
ADJ_{ij}	0.16 (1.03)	0.14 (0.85)	0.35 (2.18)	0.59 (3.72)	0.74 (4.88)
$LANG_{ij}$	0.06 (0.65)	0.34 (3.48)	0.56 (5.84)	0.80 (8.16)	0.72 (7.71)
FTA_{ij}	0.63 (3.46)	1.37 (6.64)	−0.13 (−0.73)	−0.14 (−0.95)	0.29 (2.85)
Constant	−9.38 (−20.44)	−12.17 (−26.88)	−16.23 (−35.59)	−17.09 (−40.37)	−17.94 (−49.11)
RMSE	1.4163	1.7616	1.8900	1.9919	1.9645
R^2	0.6061	0.6334	0.6446	0.6649	0.7137
No. observations	2633	4030	5421	6474	7302

t-statistics are in parentheses. The dependent variable is the (natural log of the) nominal bilateral trade flow from *i* to *j*.

unbiased coefficient estimates of β_0, β_3, β_4, β_5 and β_6 — is estimation of Eq. (2) using country-specific fixed effects.[7]

As preliminary empirical support for our claim that FTA coefficients are biased, we estimate a typical cross-section gravity equation for multiple years: 1960, 1970, 1980, 1990, and 2000. Table 1 provides cross-section coefficient estimates for these years of a typical (log-linear version of) gravity Eq. (1), estimated using the (non-zero) nominal trade flows among 96 countries identified in the Data Appendix. The data sources are standard and are provided in Section 4.[8] The five sets of estimates in Table 1 support our claims and findings in the literature that the FTA dummy's coefficient estimates are highly unstable from year to year, are often small positive values, and in some years are even *negative*.[9]

Table 2 provides estimates for the same years of gravity Eq. (2), where — following Anderson and van Wincoop (2003, Table 6) and Feenstra (2004, Ch. 5) — we use country fixed effects to account for the multilateral price terms (rather than a custom nonlinear least squares program). The FTA dummy coefficient estimates remain unstable across years and, for several years, the FTA coefficient's estimate is negative.[10] While the country fixed effects help to account for the endogeneity bias created by prices and the influence of FTAs among other

[7] Eaton and Kortum (2002), Rose and van Wincoop (2001), and Redding and Venables (2000) similarly used country fixed effects to account for multilateral price terms.

[8] The issue of zero trade flows has been treated in other papers, such as Eichengreen and Irwin (1995) and Felbermyer and Kohler (2004). However, this issue is beyond the scope of the present paper. For instance, in 2000, we have 9120 potential trade flows among 96 countries (96×95=9120), less 1573 trade flows recorded as zero and 245 observations recorded as "missing."

[9] The basic findings in Table 1 (and Table 2) are generally unaffected when we considered fixed sample sizes for all five years. First, we restricted our sample to all trading partners who had positive trade flows in 1960; for any of these pairs that had zeros in 1970, ..., 2000, we substituted ones for zeros. Second, we restricted our sample to all trading partners who had positive trade flows in 2000; for any of these pairs that had zeros in 1960, ..., 1990, we substituted ones for zeros. The coefficient estimates were materially the same across samples.

[10] The reader may also notice the upward trend (in absolute values) in non-FTA coefficient estimates with time in Tables 1 and 2. This also is an interesting issue but is beyond the scope of this paper; the upward drift (in absolute value) of the distance coefficient estimate has been studied by other researchers, cf., Felbermyer and Kohler (2004). Our focus is instead on the instability of the FTA coefficient estimates. Indeed, the FTA coefficient estimate also had the highest coefficient of variation of all the variables' coefficient estimates. For instance, using Table 2's specification, the coefficients of variation of the coefficient estimates for the non-FTA variables ranged from 0.29 to 0.31. By contrast the coefficient of variation of the FTA coefficient estimates was 2.05, six times that of the others.

Table 2
Theory-motivated cross-section gravity equations with country fixed effects

Variable	(1) 1960	(2) 1970	(3) 1980	(4) 1990	(5) 2000
ln $DIST_{ij}$	−0.68 (−16.77)	−0.89 (−21.58)	−1.28 (−31.36)	−1.30 (−31.65)	−1.46 (−35.79)
ADJ_{ij}	0.31 (2.26)	0.35 (2.38)	0.43 (2.95)	0.58 (3.93)	0.59 (4.09)
$LANG_{ij}$	0.38 (3.99)	0.84 (8.33)	0.82 (8.06)	0.98 (9.41)	0.97 (9.78)
FTA_{ij}	0.01 (0.09)	0.61 (3.27)	−1.44 (−8.65)	−1.08 (−7.30)	−0.14 (−1.36)
Constant	−14.06 (−8.25)	−12.49 (−18.66)	−14.98 (−19.37)	−16.64 (−31.88)	−12.76 (−27.18)
RMSE	1.1826	1.5025	1.6635	1.7806	1.7851
Within R^2	0.5020	0.4300	0.3857	0.3648	0.3845
No. observations	2633	4030	5421	6474	7302

t-statistics are in parentheses. The dependent variable is the (natural log of the) nominal bilateral trade flow from country i to country j divided by the product of their nominal GDPs. Coefficient estimates of country fixed effects are not reported for brevity.

countries on the trade from i to j, they do not correct for the bias introduced if countries select into FTAs. We must turn to other methods to identify unbiased FTA treatment effects, which is the purpose of this paper.[11]

3. Endogeneity bias

A standard problem in cross-section empirical work is the potential endogeneity of RHS variables. If any of the RHS variables in Eqs. (1) or (2) are correlated with the error term, ϵ_{ij}, that variable is considered econometrically "endogenous" and ordinary least squares (OLS) may yield biased and inconsistent coefficient estimates. Potential sources of endogeneity bias of RHS variables' coefficient estimates generally fall under three categories: omitted variables, simultaneity, and measurement error (see Wooldridge, 2002, pp. 50–51). While we believe all three factors may contribute potentially to endogeneity bias caused by FTA, we will argue that the most important source is omitted variables (and selection) bias. We discuss each source in turn.

3.1. Omitted variables (and selection) bias

The gravity equation has surfaced over the past four decades as the dominant empirical framework for analyzing bilateral trade flows primarily because of its strong explanatory power. Explanatory power (R^2) generally ranges from 60 to 80%. However, as Table 1 suggests, there remains considerable unobserved heterogeneity among country pairs (leaving aside the issue of multilateral price terms, for now).

In determining the potential correlation between the gravity equation's error term, ϵ_{ij}, with FTA_{ij}, one first needs to consider what determines the likelihood of an FTA between a pair of countries. Although trade economists have examined empirically for many years the determinants of tariff rates and non-tariff barrier levels across industries and across countries, virtually no empirical work has examined the determinants of FTA_{ij}. A notable exception is Baier and Bergstrand (2004a). These authors present a theoretical and empirical model of economic determinants of FTAs. Their paper finds strong cross-section empirical evidence that pairs of

[11] We emphasize that our focus here is only on an unbiased estimate of the average *treatment* effect (β_6 in Eqs. (1) and (2)), not an estimate of the full "comparative-static" effect of an FTA, which would also account for the effects of endogenous changes in the multilateral price terms, (3.1) through (3.N), cf., Anderson and van Wincoop (2003).

countries that have FTAs *tend* to share economic characteristics that their theory suggests should enhance the net economic welfare gains from an FTA for the pairs' representative consumers. For instance, two countries tend to have an FTA the larger and more similar their GDPs, the closer they are to each other but the more remote the pair is from the rest-of-the-world (ROW), and the wider (narrower) the difference in their relative factor endowments with respect to each other (ROW). But this list includes the same factors that tend to explain large trade flows. Thus, in terms of *observable* economic characteristics, countries with FTAs have "chosen well," in the sense that most country pairs with FTAs tend to have the economic characteristics associated with considerable trade and with (in theory) welfare-enhancing net trade creation from an FTA. Yet, the estimated probit functions in Baier and Bergstrand (2004a) have pseudo-R^2 values of only 70%, still leaving considerable unobserved heterogeneity.

The important question for this paper is: How is the unobserved heterogeneity in trade flow determinants associated with the likelihood of an FTA? For instance, error term ϵ_{ij} in Eqs. (1) or (2) may be representing unobservable (to the econometrician) policy-related barriers — tending to reduce trade between two countries — that are not accounted for by standard gravity equation RHS variables but may be correlated with the decision to form an FTA. As an example, suppose two countries have extensive unmeasurable domestic regulations (e.g., internal shipping regulations) that inhibit trade (causing ϵ_{ij} to be negative). The likelihood of the two countries' governments selecting into an FTA may be high if there is a large expected welfare gain from potential bilateral trade creation if the FTA deepens liberalization beyond tariff barriers into domestic regulations (and other non-tariff barriers). Thus, FTA_{ij} and the intensity of domestic regulations may be positively correlated in a cross-section of data, but the gravity equation error term ϵ_{ij} and the intensity of domestic regulations may be negatively correlated. This reason suggests that FTA_{ij} and ϵ_{ij} are negatively correlated, and the FTA coefficient will tend to be *underestimated*.

In support of this argument, numerous authors have noted that one of the major benefits of regionalism is the potential for "deeper integration." Lawrence (1996, p. xvii) distinguishes between "international policies" that deal with border barriers, such as tariffs, and "domestic policies" that are concerned with everything "behind the nation's borders, such as competition and antitrust rules, corporate governance, product standards, worker safety, regulation and supervision of financial institutions, environmental protection, tax codes...." and other national issues. The GATT and WTO have been remarkably effective in the post-WWII era reducing border barriers such as tariffs. However, these institutions have been much less effective in liberalizing the domestic policies just named. As Lawrence states it, "Once tariffs are removed, complex problems remain because of differing regulatory policies among nations" (p. 7). He argues that in many cases, FTA "agreements are also meant to achieve deeper integration of international competition and investment" (p. 7). Gilpin (2000) echos this argument: "Yet, the inability to agree on international rules or to increase international cooperation in this area has contributed to the development of both managed trade *and regional arrangements*" (p. 108; italics added). Preeg (1998) notes:

> [Free] trade agreements over time, however, have tended to include a broader and broader scope of other trade-related policies. This trend is a reflection, in part, of the fact that as border restrictions [tariffs] are reduced or eliminated, other policies become relatively more important in influencing trade flows and thus need to be assimilated in the trade relationship (p. 50).

We believe this omitted variable (selection) bias is the major source of endogeneity facing estimation of FTA effects in gravity equations using cross-section data. Moreover, the arguments above suggest that policymakers' decisions to select into an FTA are likely related to the *level* of trade (relative to its potential level), and not to recent changes in trade levels. Thus, the determinants of FTA are likely to be cross-sectional in nature.

With cross-section data, standard econometric techniques to address omitted variables (and selection) bias include estimation using instrumental variables and Heckman control functions. Alternatively, with panel data, fixed effects and first differencing can be employed to treat endogeneity bias. We discuss these various approaches in Sections 4 and 5, respectively.

3.2. Simultaneity bias

Consider the potential endogeneity bias created by simultaneity. GDP — a function of net exports — is potentially endogenous to bilateral trade flows, as suggested by the recent literature on "trade and growth," cf., Frankel and Romer (1999). Yet, some convincing reasons exist for largely ignoring potential endogeneity of incomes here. First, GDP is a function of *net* multilateral exports. Typically, net exports tend to less than 5% of a country's GDP (in absolute terms). Moreover, while GDP is related to net exports, GDP's connection to gross exports is much less direct. Second, the gravity equation relates *bilateral* trade flows to countries' incomes. Trade between any pair of countries tends to be a very small share of any country's multilateral exports, much less its GDP. Note that the literature on "trade and growth" examines empirically the effects of multilateral — not bilateral — trade on GDP, cf., Frankel and Romer (1999). Third, Frankel (1997) and others have previously accounted for the potential endogeneity of national incomes econometrically in typical gravity equations using instrumental variables (IV) including labor forces and stocks of human and physical capital. Frankel (1997) reported that coefficient estimates in gravity equations change insignificantly using these IV techniques and concluded "Evidently, the endogeneity of income makes little difference" (p. 135). Nevertheless, we will still also account for the potential endogeneity of GDPs using a regression specification such as Eq. (2) with GDPs on the left hand side (LHS).

Of the remaining RHS variables in Eq. (1), only FTA seems potentially endogenous. It is plausible to treat DIST and ADJ as exogenous. Moreover, the presence or absence of a common language (LANG) may reasonably be treated in cross-section as exogenous also.

Yet, as discussed earlier, there exists a large empirical literature in international trade on the effects of multilateral tariff and nontariff barriers on multilateral trade volumes, and the simultaneous effects of these trade volumes on multilateral barriers using cross-industry and cross-country data for particular years, c.f. Trefler (1993) and Lee and Swagel (1997). Simultaneity may be an issue for FTA in cross-section gravity equations, motivated as in these two studies. For example, holding constant typical gravity equation RHS variables (GDPs, DIST, ADJ, LANG), two countries (say, the United States and China) that possibly trade more than their "natural" level, as predicted by a typical gravity equation, may create political pressures to avoid trade liberalization or possibly raise trade barriers. This would cause a negative simultaneity bias in the FTA coefficient estimate. On the other hand, the governments of two countries that trade more than their gravity-equation-suggested "natural" level might be induced to form an FTA because there might potentially be less "trade diversion" due to their extensive trading relationship, suggesting a positive simultaneity bias. However, since the decisions to select into FTAs are likely influenced by the levels of trade relative to "natural" levels, recent *changes* in trade levels are not likely to influence FTA formations.

To address this issue, the natural inclination is to estimate a system of simultaneous equations treating bilateral trade and FTAs as endogenous variables using instrumental variables, following in the spirit of Trefler (1993) and Lee and Swagel (1997). We will discuss some of these efforts in Section 4.

3.3. Measurement error bias

Finally, measurement error in an explanatory variable, such as an FTA dummy, is generally associated with negative bias (in absolute terms) in the variable's coefficient. For instance, with the classical "errors-in-variables" assumption, the 0–1 FTA dummy variable (FTA) would be correlated positively with the measurement error (ς) if the true trade-policy variable (say, the tariff rate, t) was assumed uncorrelated with ς ($\varsigma = \mathrm{FTA} - t$). In Eq. (1)'s context, the correlation between FTA and the error term ($\epsilon - \beta_6\,\varsigma$) would be negative, leading to the classical "attenuation bias" of FTA's coefficient estimate toward zero.[12] We believe that this may be part of the reason — but neither the entire, nor even the most important, reason — FTA coefficient estimates have been underestimated.

Of course, the best method for eliminating this bias is construction of a continuous variable that would more accurately measure the degree of trade liberalization from various FTAs. If FTAs only eliminated bilateral tariff rates, one would ideally measure this liberalization with a change in the ad valorem tariff rate (for which data is poor). However, FTAs liberalize trade well beyond the elimination of tariffs. Calculation of such measures is beyond the scope of this particular study, but is a useful direction for future research. Our goal rather is to determine reliable estimates of the *treatment effect* of an FTA, similar to the 0–1 variable representing program participation in labor econometrics. Thus, we constrain our study to estimate more accurately the ex post effect of an FTA dummy on trade flows, as has been employed in the gravity equation literature for over four decades.[13]

4. Treatment effects in cross-section models of trade flows and free trade agreements

We briefly address conventional cross-section instrumental-variables (IV) and control-function approaches to address the FTA endogeneity bias associated with omitted variables and selection. We then discuss some previous studies using these techniques and summarize their mixed findings.

4.1. Cross-section econometric techniques

Methodological issues regarding the cross-section estimation of the partial effects of an endogenous binary variable (such as FTA) on a continuous endogenous variable (such as trade flows) fall under the "treatment effect" literature in econometrics. The average treatment effect refers to the notion that the trade flow between two countries will differ depending upon whether the countries share an FTA or not. The fundamental econometric dilemma is that one can observe only one situation or the other. An excellent summary of developments in the treatment–effect

[12] Even without the classical errors-in-variable assumption, the correlation between FTA and ϵ is likely negative. Suppose the true trade policy variable is the bilateral tariff rate, t, where $t>0$. If an FTA exists, $\mathrm{FTA}=1$, $t=0$, and ς consequently equals 1. If no FTA exists, $\mathrm{FTA}=0$, $t>0$, and consequently $\varsigma<0$. Thus, FTA and ς are positively correlated.

[13] Another interesting direction suggested by one referee is to also examine the treatment effects by individual agreements. However, to limit the scope of this paper, this useful direction is left for future research.

literature is in Heckman (2001) and Wooldridge (2002, Ch. 18); we refer the reader to these papers for a more comprehensive discussion.

The average treatment effect (ATE) of an FTA between a country pair is defined as:

$$\mathrm{ATE}(\boldsymbol{q}) \equiv E(x_1 - x_0 | \boldsymbol{q}, \mathrm{FTA}) \tag{4}$$

where x_1 (x_0) denotes the logarithm of the trade flow from country i to country j with (without) the FTA, E denotes the expectation operator, and observation subscripts (ij) are omitted for simplicity. We assume that the level of trade between two countries depends upon the array of exogenous "covariates" other than the treatment — the standard set of gravity equation variables in Eq. (1) in levels or log levels as appropriate, excluding FTA — which we denote by $\boldsymbol{q}$ (GDPs, DIST, ADJ, LANG). Initially, we assume that observations of x are independently and identically distributed across country pairs; this assumption ensures that the treatment of one pair does not affect another pair's trade flow.

Consistent estimation of the average treatment effect depends upon assumptions made about relationships among the variables. Of course, in reality we do not observe both x_1 and x_0. Since one can only observe a trade flow in the presence *or* absence of an FTA, we define the observed outcome (x) for a country pair as:

$$x \equiv (\mathrm{FTA})x_1 + (1 - \mathrm{FTA})x_0 \tag{5}$$

where FTA = 1 if an FTA exists between the pair, and 0 otherwise. If we assume trade flows x_0 and x_1 have the standard linear form as in a gravity equation, then:

$$x_0 = \mu_0 + \boldsymbol{\beta}'\boldsymbol{q} + \varepsilon_0 \tag{6}$$

$$x_1 = \mu_1 + \boldsymbol{\beta}'\boldsymbol{q} + \varepsilon_1 \tag{7}$$

Substituting Eqs. (6) and (7) in Eq. (5) yields:

$$x = \mu_0 + \boldsymbol{\beta}'\boldsymbol{q} + \alpha\mathrm{FTA} + \varepsilon_0 + \mathrm{FTA}(\varepsilon_1 - \varepsilon_0) \tag{8}$$

where $\alpha = \mu_1 - \mu_0$ corresponds to the average treatment effect.

Consistent estimation of the parameters in Eq. (8) depends upon correlations between the variables and the error terms. Specifically, consistent estimation of α, $\boldsymbol{\beta}$, and μ_0 depends on the correlation of: (i) FTA with error term ϵ_0, and (ii) FTA with differences in unobservables for partners with FTAs versus partners without FTAs ($\epsilon_1 - \epsilon_0$). The former is the correlation associated with omitted variables bias, and the latter with selection bias. When FTA is uncorrelated with both factors, the parameters are estimated consistently using OLS.

Consider first the case where FTA is correlated with ϵ_0, but $\epsilon_1 - \epsilon_0 = 0$ (for all country pairs). Wooldridge (2002) suggests an efficient IV estimator. First, assume the probability of FTA (P(FTA)) can be estimated by a known parametric form (such as probit) such that $P(\mathrm{FTA}=1|\boldsymbol{q}, \boldsymbol{z}) = \boldsymbol{\Phi}(\pi_0 + \boldsymbol{\pi}_1'\boldsymbol{q} + \boldsymbol{\pi}_2'\boldsymbol{z})$ where $P(\mathrm{FTA}=1|\boldsymbol{q}, \boldsymbol{z}) \neq P(\mathrm{FTA}=1|\boldsymbol{q})$ and $\boldsymbol{z}$ is a set of exogenous variables (not in $\boldsymbol{q}$).[14] Second, assume the variance of ϵ_0 is a constant. This suggests a multi-step IV method with the following steps: (i) estimate a binary response model, such as probit, $P(\mathrm{FTA}=1|\boldsymbol{q}, \boldsymbol{z}) = \boldsymbol{\Phi}(\pi_0 + \pi_1'\boldsymbol{q} + \boldsymbol{\pi}_2'\boldsymbol{z})$ by maximum likelihood to generate predicted probabilities $\boldsymbol{\Phi}^{\mathrm{P}}$; and (ii) estimate Eq.

[14] If a probit function, then $\boldsymbol{\Phi}(\ )$ is the standard normal cumulative distribution function.

(8) by IV, using instruments 1, $\boldsymbol{\Phi}^{\mathrm{P}}$, $\boldsymbol{q}$ and $\boldsymbol{z}$. The IV estimator is consistent, asymptotically efficient, and the usual 2SLS standard errors and test statistics are asymptotically valid.[15]

If FTA is correlated with ϵ_0 and $\epsilon_1 - \epsilon_0$, Heckman (1997) suggests that IV estimation only identifies the parameters consistently when economic agents' decisions to select into a program are unrelated to (or ignore) unobservable factors influencing the outcome. However, as discussed earlier, many trade-policy analysts have noted that policies tending to inhibit trade, such as nontariff barriers and domestic regulations, may be one of the main reasons governments have selected into FTAs. The intent of forming an FTA is that this agreement will lead to bilateral reductions in domestic barriers that multilateral agreements have been unable to attain. In this context, IV estimation will not yield consistent estimates in the presence of selection bias; Heckman's procedure controls for selection.

As discussed earlier, most cross-section analyses using the gravity equation including dummy variables for trade agreements have ignored the potential endogeneity bias just discussed. The introduction listed several of a myriad of gravity equation studies that have tried to infer inappropriately the average treatment effect of a trade agreement on bilateral trade ignoring this bias. The fragility of such ATEs reported in the analysis of Ghosh and Yamarik (2004) tends to confirm the likely presence of endogeneity bias.

4.2. Cross-section empirical attempts to adjust for endogeneity bias

As just established, given a cross-section the key to estimating a consistent ATE of FTAs on trade is, first, finding a probit function that predicts FTAs and, second, finding a set of suitable instruments for z that are uncorrelated with the gravity equation error term. Only three papers have attempted to adjust for the potential endogeneity of FTAs using IV or control-function techniques with cross-section data, Baier and Bergstrand (2002, 2004b) and Magee (2003). We discuss each of their findings in turn.

The starting point for all three studies is Baier and Bergstrand (2004a), which evaluated the probability of pairs of countries having FTAs using a probit function for a single year. As discussed in Section 3, Baier and Bergstrand (2004a) was the first systematic empirical analysis of economic determinants of the likelihood of FTAs between pairs of countries using a cross-section qualitative-choice model. Their empirical model correctly predicted *85%* of 286 FTAs existing in 1996 among 1431 country pairs and *97%* of the remaining 1145 pairs, based upon economic and geographic characteristics such as economic size and similarity of the country pair, differences in capital–labor ratios, and bilateral distance of the pair and their remoteness from the rest of the world (ROW).

Baier and Bergstrand (2002) followed the Wooldridge procedure described in Section 4.1, essentially augmenting the probit function in Baier and Bergstrand (2004a) to include a set of instruments that ideally would be correlated with the probability of an FTA between a country pair but uncorrelated with (unobservables causing) their bilateral trade. Critical to the methodology is the selection of instruments. The study first used relative capital–labor ratios, relative factor-endowment differences with the ROW, and an ad hoc measure of remoteness of continental FTA partners as instruments using a cross-section of bilateral trade flows among 53 countries for year

[15] Wooldridge (2002) refers to this as a two-step estimator. However, for clarity, we refer to three steps. The first stage is the estimation of the predicted probabilities, $\boldsymbol{\Phi}^{P}$. The second stage is a linear regression of FTA on a constant, $\boldsymbol{\Phi}^{P}$, $\boldsymbol{q}$ and $\boldsymbol{z}$. The third stage is estimation of the gravity equation substituting the predicted values from the second-stage regression for FTA.

1996. While these instruments "worked" in the sense of quadrupling the average treatment effect of an FTA, two concerns arose. First, the Wooldridge estimation procedure precluded a test of over-identifying restrictions to provide empirical support that the instruments were, in fact, "exogenous" to the gravity equation error terms. More importantly, trade gravity equations have often included measures of remoteness and capital–labor ratio differences and these factors have had statistically significant effects, eroding confidence that these instruments were uncorrelated with gravity equation errors. Baier and Bergstrand (2002) also considered various "political" variables as instruments. However, the same concern arises: many of the "political" variables have been correlated in previous studies with trade flows. Despite estimating an ATE of about 90%, the instability of estimated treatment effects — ranging from increasing trade between 40 to 200% — further diminished confidence in the estimates.[16]

Baier and Bergstrand (2004b) demonstrated clearly that ATE estimates using IV or control-function techniques are quite unstable. This paper expanded the data set to trade flows and FTAs among 96 countries for multiple years. Using alternative specifications — with and without political variables, with and without fixed effects (as appropriate) — this study found that ATE estimates ranged between a decline in trade of 92% to an increase in trade of 1100%. In most specifications, the test for overidentifying restrictions rejected the null hypothesis that the instruments were exogenous. In the only two specifications where this null hypothesis could not be rejected (using a linear probability model in the first stage), the two ATE estimates were 0.41 (statistically insignificant) and −3.97 (statistically significant), varying according to the instruments included.

Magee (2003) is also one of the first papers to adjust for endogeneity of FTAs using instrumental variables. Magee (2003) viewed the relationship between trade and FTA as a simultaneous-equations system.[17] Magee uses 2SLS to estimate the effect of endogenous FTAs on trade flows, and finds similarly a range of large positive to large negative effects of FTAs on trade flows. Magee's study faces the same limitations as Baier and Bergstrand (2002, 2004b). Several of the instruments in the FTA probit equation to identify the trade-flow equation are likely correlated with the gravity equation error terms. For instance, Magee includes an index of democracies, GDP similarities, intra-industry trade indices, trade surpluses and relative-factor-endowment differences. All these variables are likely correlated in cross-section with (unobservables causing) trade flows.

We conclude from previous cross-section studies that IV estimation is not a reliable method for addressing the endogeneity bias of the FTA binary variable in a gravity equation, despite trying a wide array of economic and political instrumental variables. An alternative method for estimating the ATE of FTAs uses Heckman's control-function approach. We have estimated similar specifications using this alternative approach with qualitatively similar findings; the control-function approach does not solve the endogeneity bias issue either. The likely problem is this: the vast number of variables that are correlated cross-sectionally with the probability of having an FTA are *also* correlated cross-sectionally with trade flows, preventing elimination of the endogeneity bias using cross-section techniques.

Magee (2003) concluded that "we should be cautious in using gravity equation estimates to draw strong conclusions about the effect of PTA formation on trade." We agree with his conclusion for *cross-sectional data*. However, in the remainder of this paper, we argue that one can draw strong and reliable inferences about the ATE of FTAs using the gravity equation applied to *panel data*.

[16] The results using Heckman control functions were not materially different.

[17] See Baier and Bergstrand (2002, 2004b) and Maddala (1983, p. 118) for a critique of this approach in the context of a binary endogenous variable.

5. FTA treatment effects using panel data

As most standard econometrics textbooks now suggest, a ready alternative to cross-section estimation of treatment effects in the presence of unobserved time-invariant heterogeneity is the use of panel data (cf., Wooldridge, 2000). Having constructed a panel (for every five years) from 1960–2000 of the bilateral trade flows, bilateral trade agreements, and standard gravity equation covariates among 96 potential trading partners, we now pursue this approach. Three main alternative techniques exist for addressing the issue: estimating the panel with random effects, estimation with fixed effects, or differencing the data and using OLS. In Section 5.1, we describe our data set. In Section 5.2, we address the choice between random versus fixed effects, and then between fixed effects versus first differencing. In Section 5.3, we estimate the atheoretical gravity model (ignoring multilateral price terms) using fixed effects. In Section 5.4, we estimate the model again using fixed effects, but address three issues. First, we allow for the theoretically-motivated multilateral price terms and unit income elasticities. Second, we allow for the "phasing-in" of FTAs. Third, we test for "strict exogeneity." In Section 5.5, we provide the results of a robustness analysis using differenced data.

5.1. Data

We describe briefly the data used for the gravity equations. Nominal bilateral trade flows are from the International Monetary Fund's *Direction of Trade Statistics* for the years 1960, 1965,…, 2000 for 96 potential trading partners (zero trade flows are excluded); these data are scaled by exporter GDP deflators to generate real trade flows for the panel analysis. Nominal GDPs are from the World Bank's World Development Indicators (2003); these are scaled by GDP deflators to create real GDPs for the panel analysis. Bilateral distances were compiled using the *CIA Factbook* for longitudes and latitudes of economic centers to calculate the great circle distances. The language and adjacency dummy variables were compiled also from the *CIA Factbook.* The FTA dummy variable was calculated using appendices in Lawrence (1996) and Frankel (1997), various websites, and FTAs notified to the GATT/WTO under GATT Articles XXIV or the Enabling Clause for developing economies; we included only full (no partial) FTAs and customs unions. Table 3 lists the trade agreements used and sources. The Data Appendix lists the countries used.[18]

5.2. Alternative panel methodologies

5.2.1. Fixed versus random effects

Our panel estimation applies fixed effects rather than random effects for two reasons, the first on conceptual grounds and the second on empirical grounds. First, as addressed in Section 2, we believe the source of endogeneity bias in the gravity equation is unobserved time-invariant heterogeneity. In economic terms, we believe there are unobserved time-invariant bilateral variables — termed w_{ij} — influencing simultaneously the presence of an FTA and the volume of trade. Because these variables are likely correlated with FTA_{ij}, they are best controlled for using bilateral "fixed effects," as this approach allows for arbitrary correlations of w_{ij} with FTA_{ij}. By contrast, under "random effects" one assumes zero correlation between unobservables w_{ij} with FTA_{ij}, which seems less plausible.

Second, there have been recent econometric evaluations of the gravity equation with panel data using the Hausman Test to test for fixed versus random effects. For example, Egger (2000) finds

[18] The data set is available at the authors' websites (http://www.nd.edu/~jbergstr and http://people.clemson.edu/~sbaier).

Table 3
Free trade agreements

European Union, or EU (1958): Belgium–Luxembourg, France, Italy, Germany, Netherlands, Denmark (1973), Ireland (1973), United Kingdom (1973), Greece (1981), Portugal (1986), Spain (1986), Austria (1995), Finland (1995), Sweden (1995)
The Customs Union of West African States (1959): Burkina Faso, Mali, Mauritania, Niger, Senegal
European Free Trade Association, or EFTA (1960): Austria (until 1995), Denmark (until 1973), Finland (1986–1995), Norway, Portugal (until 1986), Sweden (until 1995), Switzerland, United Kingdom (until 1973)
Latin American Free Trade Agreement/Latin American Integration Agreement, or LAFTA/LAIA (1961–1979,1993–): Argentina, Bolivia, Brazil, Chile, Ecuador, Mexico, Paraguay, Peru, Uruguay, Venezuela (became inoperative during 1980–1990, but reinitiated in 1993)
African Common Market (1963): Algeria, Egypt, Ghana, Morocco
Central American Common Market (1961–1975, 1993–present): El Salvador, Guatemala, Honduras, Nicaragua, Costa Rica (1965)
Economic Customs Union of the Central African States (1966): Cameroon, Congo, Gabon
Carribean Community, or CARICOM (1968): Jamaica, Trinidad and Tobago, Guyana (1995)
EU–EFTA Agreement/European Economic Area (1973/1994)
Australia–New Zealand Closer Economic Relations (1983)
US–Israel (1985)
US–Canada (1989)
EFTA–Israel (1993)
Central Europe Free Trade Agreement, or CEFTA (1993): Hungary, Poland, Romania (1997), Bulgaria (1998)
EFTA–Bulgaria (1993)
EFTA–Hungary (1993)
EFTA–Poland (1993)
EFTA–Romania (1993)
EU–Hungary (1994)
EU–Poland (1994)
North American Free Trade Agreement, or NAFTA (1994): Canada, Mexico, United States
Bolivia–Mexico (1995)
Costa Rica–Mexico (1995)
EU–Bulgaria (1995)
EU–Romania (1995)
Group of Three (1995): Columbia, Mexico, Venezuela
Mercado Comun del Sur, or Mercosur (1991): Argentina, Brazil, Paraguay, Uruguay (formed in 1991 and a free trade area in 1995)
Andean Community (1993): Bolivia, Columbia, Ecuador, Peru, Venezuela, Peru (1997)
Mercosur–Chile (1996)
Mercosur–Bolivia (1996)
Canada–Chile (1997)
Canada–Israel (1997)
Association of Southeast Asian Nations, or ASEAN (1998): Indonesia, Philippines, Singapore, Thailand (effective on 80% of merchandise trade in 1998)
CARICOM–Dominican Republic (1998)
Hungary–Turkey (1998)
Hungary–Israel (1998)
India–Sri Lanka (1998)
Israel–Turkey (1998)
Mexico–Nicaragua (1998)
Romania–Turkey (1998)
Poland–Israel (1998)
Romania–Turkey (1998)
Mexico–Chile (1999)
Common Market for Eastern and Southern Africa (2000): Egypt, Kenya, Madagascar, Malawi, Mauritius, Sudan, Zimbabwe, Zambia

(continued on next page)

Table 3 (*continued*)

EU–Israel Agreement (2000)
EU–Mexico (2000)
Poland–Turkey (2000)
Mexico–Guatemala (2000)
Mexico–Honduras (2000)
Mexico–Israel (2000)
Mexico–El Salvador (2000)
New Zealand–Singapore (2000)

Countries listed in agreements only include those in our sample of 96 countries listed in the Data Appendix. Agreements are listed in chronological order of date of entry into force. Years in parentheses denote date of entry, except where noted otherwise.

Sources:
http://www.wto.org/english/tratop_e/region_e/summary_e.xls.
http://europa.eu.int/comm/enlargement/pas/europe_agr.htm.
http://www.comunidadandina.org/ingles/union.htm.
http://www.nafinsa.com/finsafreetrade.htm.
http://www.sice.oas.org/default.asp.
Frankel, Jeffrey A. Regional trading blocs. Institute for International Economics (1997).
Lawrence, Robert Z. Regionalism, multilateralism, and deeper integration. The Brookings Institution (1996).

overwhelming evidence for the rejection of a random-effects gravity model relative to a fixed-effects gravity model, using either bilateral-pair or country-specific fixed effects.

5.2.2. *Fixed effects versus first differencing*

Standard econometric discussions of treating endogeneity bias using panel data focus on a choice between estimation using fixed effects versus using first-differenced data, cf., Wooldridge (2002, Ch. 10). As Wooldridge notes, when the number of time periods (T) exceeds two, the fixed-effects estimator is more efficient under the assumption of serially uncorrelated error terms. The first-differencing estimator is more efficient (when $T>2$) under the assumption that the error term ϵ_{ijt} follows a random walk (i.e., that the difference in the error terms, $\epsilon_{ijt}-\epsilon_{ij,t-1}$, is white noise).[19]

First-differencing the panel data yields some potential advantages over fixed effects. First, it is quite plausible that the unobserved heterogeneity in trade flows, ϵ_{ijt}, is correlated over time. In light of discussion in Section 3, unobserved factors influencing the likelihood of an FTA (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 fixed effects 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 fixed effects is equivalent to differencing data around the *mean* (in our sample, 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 fixed effects. First-differencing yields data that deviates from the previous period of our panel, and thus is closer to a unit-root process. In the following, we use fixed effects in Sections 5.3 and 5.4, and for robustness use differenced data in Section 5.5.

[19] When the number of time periods is limited to two ($T=2$), estimation with fixed effects and first-differencing produce *identical* estimates and inferences; moreover, first-differencing is easier. When $T>2$, the choice depends upon the assumption the researcher makes about the error term ϵ_{ijt}.

5.3. Fixed-effects estimation of an atheoretical gravity equation ignoring multilateral price terms

In a panel context, Eq. (1) can be expressed as:

$$\begin{aligned}\ln X_{ijt} = \beta_0 + \beta_1(\ln\mathrm{RGDP}_{jt}) + \beta_2(\ln\mathrm{RGDP}_{jt}) + \beta_3(\ln\mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{ij}) \\ + \beta_5(\mathrm{LANG}_{ij}) + \beta_6(\mathrm{FTA}_{ijt}) + \varepsilon_{ijt}\end{aligned} \tag{9}$$

Table 4 provides the empirical results of estimating gravity Eq. (9) using a panel of real trade flows (X_{ijt}), real GDPs (RGDP_{it}, RGDP_{jt}) and FTA dummies (FTA_{ijt}), and using alternative specifications with and without bilateral fixed effects and time dummies. Column (1) provides the baseline gravity equation without any fixed effects or time dummies for all nine years. Exporter and importer (real) GDPs have coefficients close to unity, distance has a traditional coefficient estimate of −1, and the adjacency and language dummies have typical coefficient estimates. The ATE for FTA of 0.13 is economically small, consistent with earlier findings in cross-section. Column (2) provides the empirical results including a time dummy, where (for brevity) we omit reporting the (statistically significant) coefficient estimates for these dummy variables. The inclusion of time dummies increases the ATE for FTA slightly from 0.13 to 0.27. However, time dummies do not adjust for the endogeneity of FTAs.

Adjusting for unobserved time-invariant heterogeneity using bilateral fixed effects has a notable impact on the results. Column (3) provides results including bilateral fixed effects. The ATE for FTA (0.51) is now almost *quadruple* the coefficient estimated using OLS (0.13). Column (4) provides results using both bilateral fixed effects and year dummies. In this specification, the ATE of a free trade agreement is 0.68. This quantitative estimate suggests that the average treatment effect of the presence of a free trade agreement is to *double* trade between country pairs ($e^{0.68}=1.97$, or 97% increase). This is seven times the effect estimated using OLS ($e^{0.13}=1.14$, or 14% increase).[20]

Although the introduction of bilateral fixed effects and time dummies has an economically and statistically significant effect on raising the FTA average treatment effect, this specification suffers ex ante from ignoring multilateral price terms, as suggested by recent theoretical developments. In the next section, we account for such terms, as well as the potential influence of "phasing-in" agreements.

5.4. Fixed-effects estimation of a theoretically-motivated gravity equation with phased-in agreements

In this section, we consider three modifications to the previous specification. In Section 5.4.1, we include country-and-time effects to account for the theoretically-motivated multilateral price terms and consider unit income elasticities. In Section 5.4.2, we account for the fact that all FTAs are "phased-in" over time, typically over five-to-ten years, and for the possibility that the change in two members' terms of trade from formation of an FTA may have a lagged impact on their

[20] The only other published studies that have estimated the ATE of an FTA using a panel of data spanning as many years and at least 100 countries are Rose (2004) and Tomz, Goldstein and Rivers (2004). Using fixed effects, Rose found an ATE of $e^{0.94}$or 156%. However, using a classification of formal and informal GATT members, Tomz, Goldstein and Rivers (2004) estimate an ATE for FTAs (with fixed effects) of only $e^{0.76}$or 114%. Cheng and Wall (2002) used bilateral fixed effects in a four-year panel of trade among approximately only 30 high-income countries in the context of a traditional gravity equation ignoring multilateral price terms.

Table 4
Panel gravity equations in levels using various specifications

Variable	(1) No fixed or time effects	(2) With time effects	(3) With bilateral fixed effects	(4) With time and bilateral fixed effects
ln $RGDP_i$	0.95 (217.50)	0.97 (230.98)	0.71 (34.54)	1.27 (47.16)
ln $RGDP_j$	0.94 (224.99)	0.97 (235.43)	0.58 (26.57)	1.22 (41.60)
ln $DIST_{ij}$	−1.03 (−79.09)	−1.01 (−78.60)		
ADJ_{ij}	0.41 (8.23)	0.38 (7.28)		
$LANG_{ij}$	0.63 (19.06)	0.58 (17.73)		
FTA_{ij}	0.13 (3.73)	0.27 (7.19)	0.51 (10.74)	0.68 (14.27)
RMSE	1.9270	1.8601		
Overall R^2	0.6575	0.6809		
Within R^2			0.2036	0.2268
No. observations	47,081	47,081	47,081	47,081

t-statistics are in parentheses. The dependent variable is the (natural log of the) real bilateral trade flow from *i* to *j*. Coefficient estimates for various fixed/time effects are not reported for brevity.

bilateral trade. In Section 5.4.3, we address "strict exogeneity" issues; we test for the possibility of reverse causality by addressing the effect of *future* FTA dummies on current trade flows.

5.4.1. Accounting for multilateral price terms and unit income elasticities

While the results in the previous section are encouraging, the gravity equation suggested by recent formal theoretical developments — summarized in the system of Eqs. (2), (3.1), …, (3.N) in Section 2 — suggests that one needs to account for the multilateral price variables and to scale the LHS trade flow variable by real GDPs. None of the four specifications in Table 4 accounts for these two elements. First, accounting for the multilateral price variables in a panel context suggests estimating:

$$\begin{aligned}\ln X_{ijt} = {} & \beta_0 + \beta_1(\ln \mathrm{RGDP}_{it}) + \beta_2(\ln \mathrm{RGDP}_{jt}) + \beta_3(\ln \mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{ij}) \\ & + \beta_5(\mathrm{LANG}_{ij}) + \beta_6(\mathrm{FTA}_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \varepsilon_{ijt}\end{aligned} \tag{10}$$

Furthermore, scaling the LHS variable by the product of real GDPs suggests estimating:

$$\begin{aligned}\ln[X_{ijt}/(\mathrm{RGDP}_{it}\mathrm{RGDP}_{jt})] = {} & \beta_0 + \beta_3(\ln \mathrm{DIST}_{ij}) + \beta_4(\mathrm{ADJ}_{ij}) + \beta_5(\mathrm{LANG}_{ij}) \\ & + \beta_6(\mathrm{FTA}_{ijt}) - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \varepsilon_{ijt}\end{aligned} \tag{11}$$

In a panel setting, the multilateral price variables would be *time varying*, and consequently the results in specifications (1)–(4) in Table 4 may suffer from an omitted variables bias as a result of ignoring these time-varying terms — a dilemma that cannot be resolved by the use of bilateral fixed effects using the panel data in its current form.[21] Moreover, the theoretical model in Eqs. (2), (3.1),…, (3.N) suggests that the coefficient estimates for the real GDP variables should be unity, even though using bilateral fixed effects in specifications (3) and (4) suggests income elasticities are significantly different from unity.

[21] Random effects estimation would not be of any use either, as theory suggests that the multilateral price terms and the FTA variable would be correlated.

We first estimate Eq. (10) using bilateral (*ij*) fixed effects to account for variation in DIST, ADJ, and LANG along with country-and-time (*it*, *jt*) effects to account for variation in real GDPs and the multilateral price terms. In the context of the theory (though ignoring the restriction of unitary income elasticities), this should generate an unbiased estimate of β_6. Column (1) in Table 5 provides the results of estimating this equation using bilateral fixed effects and the country-and-time effects. As indicated, the FTA coefficient estimate of 0.46 suggests that an FTA increases trade by about 58% on average.

We can impose the restriction of unitary income elasticities implied by theoretical Eq. (2) and estimate Eq. (11) using bilateral-pair (*ij*) fixed effects along with the country-and-time (*it*, *jt*) effects. Here, the country-and-time (*it*, *jt*) effects account explicitly for the time-varying multilateral price terms. However, on net, scaling or not scaling real trade flows by real GDPs will not matter for estimating the ATE. In log-linear form, the variation in the logs of real GDPs is captured by the country-and-time effects, and only the estimates of the intercept and the country-and-time effects' coefficients change; the FTA coefficient estimate is unaffected. Column (2) in Table 5 provides the results of estimating this equation. The results using "trade shares" (real trade flows scaled by the product of real GDPs) are virtually identical to those for Eq. (10). Thus, imposing the unitary income elasticities restrictions has no impact on the FTA coefficient estimate. The FTA coefficient estimate of 0.46 implies again that an FTA increases trade by a cumulative amount of about 58%. In the remainder of the results, we use the real trade flow for the LHS variable; the FTA coefficient estimates are identical using trade shares instead (and are available on request).

The average treatment effect of an FTA accounting for the theoretically-motivated country-and-time effects, 0.46, is considerably smaller than the effect estimated in column (4) of Table 4, 0.68. However, although the estimated FTA effects in columns (1) and (2) of Table 5 are smaller, they may not reflect fully the cumulative effects on trade of an FTA.

5.4.2. Accounting for "phased-in" FTAs and lagged terms-of-trade effects

In this section, we introduce lagged effects of FTA on trade. The economic motivation for including lagged changes stems partly from the institutional nature of virtually all FTAs. The 0–1 FTA_{ijt} variable was constructed using the "Date of Entry into Force" of the agreement, as best surmised by scrutinizing multiple data sources provided earlier. However, virtually every FTA is

Table 5
Panel gravity equations with bilateral fixed and country-and-time effects[a]

Variable	(1)	(2)	(3)	(4)	(5)
$FTA_{ij,t}$	0.46** (9.07)	0.46** (9.06)	0.29** (4.95)	0.28** (4.66)	0.35** (4.20)
$FTA_{ij,t-1}$			0.38** (5.62)	0.27** (3.30)	0.16* (1.64)
$FTA_{ij,t-2}$				0.21** (2.60)	0.17* (1.87)
$FTA_{ij,t+1}$					−0.04 (−0.62)
Constant	8.85** (151.71)	−24.59** (−429.81)	9.70** (147.93)	10.06** (124.57)	9.98**(93.20)
Total ATE[a]	0.46	0.46	0.67	0.76	0.68
Within R^2	0.3102	0.1891	0.3044	0.2750	0.2516
No. observations	47,081	47,081	36,563	34,105	27,575

t-statistics are in parentheses. The dependent variable for specifications (1), (3), (4), and (5) is the (natural log of the) real bilateral trade flow; the dependent variable for specification (2) is the (natural log of the) real bilateral trade flow divided by the product of the real GDPs. * (**) denotes statistical significance at 5 (1) percent level in one-tailed *t*-test. Coefficient estimates for bilateral fixed and country-and-time effects are not reported for brevity.

[a] Total ATE is the sum of the statistically-significant FTA coefficient estimates.

"phased-in," typically over 10 years. For instance, the original EEC agreement of 1958 had a 10-year phase-in period. NAFTA had a similar 10-year provision. Thus, the entire economic (treatment) effect cannot be captured fully in the concurrent year only. It is reasonable to expect an FTA entered into "legally" in 1990 to not come into economic effect fully until 2000. Thus, it is reasonable to include one or two lagged levels of the FTA dummy ($FTA_{ij,t-1}$ and/or $FTA_{ij,t-2}$).

Moreover, economic effects of an FTA include altering the terms of trade. However, as is well known from a large literature in international economics, terms-of-trade changes tend to have lagged effects on trade volumes. Thus, it is reasonable to assume that an FTA which enters into force in 1960, and which is even fully "phased-in" by 1965, might still have an effect on trade flows in 1970.

The results in columns (3) and (4) in Table 5 reveal that FTA has statistically significant lagged effects on trade flows. Moreover, the coefficient estimates have economically plausible values, balanced across periods. The cumulative ATE with one lag is 0.65; with two lags, the total ATE is 0.76. We also experimented with adding a third lag; however, estimated coefficients of the third lag were statistically insignificant. The economic interpretation of the ATE of 0.76 is that — after 10 years — formation of an FTA increases the level of trade by *114%*.

5.4.3. Strict exogeneity

The introductory quote in this paper suggests questioning whether traditional cross-section estimates of FTA on trade can be interpreted as FTAs causing trade *or* trade causing FTAs. The results of previous sections suggest that — after accounting for endogeneity using panel data — one can find economically significant ATEs for FTA. However, to confirm that there are no "feedback effects" from trade changes to FTA changes, we run one more regression using the fixed-effects specification.[22] Wooldridge (2002, p. 285) suggests that it is easy to test for the "strict exogeneity" of FTAs in our context. To do this, we add a *future* level of FTA to the regression model. In the panel context here, if FTA changes are strictly exogenous to trade flow changes, $FTA_{ij,t+1}$ should be uncorrelated with the concurrent trade flow. The results in column (5) of Table 5 confirm this. The effect of $FTA_{ij,t+1}$ on the trade flow is economically small and not significantly different from zero. Moreover, the coefficient estimate of -0.04 suggests — if anything — that firms delay trade temporarily in anticipation of an impending agreement.

5.5. First-differenced panel gravity equation estimates

As discussed in Section 5.2, for two econometric reasons we may expect first-differenced data to provide better estimates of the average treatment effect; at worst, differenced data provide an evaluation of the robustness of previous estimates. In the context of differenced panel data, the potential omitted variables bias created by time-varying multilateral price terms for each country would require again country-and-time effects to obtain consistent estimates of FTA's ATE. As before, with country-and-time effects the coefficient estimates of the FTA treatment effects are insensitive to the real bilateral trade flow being scaled or not scaled by real GDPs; for consistency to earlier results, we present those for the flows (the virtually identical results are available on

[22] An empirical finding that trade leads an FTA need not even imply that trade "causes" an FTA. Trade may increase in anticipation of an FTA as infrastructure and delivery systems involving sunk costs are redirected, cf., McLaren (1997). Alternatively, trade may decrease — be delayed — in anticipation of the benefits of an FTA.

Table 6
First-differenced panel gravity equations with country-and-time effects[a]

Variable	(1)	(2)	(3)	(4)
$\text{dFTA}_{ij,t-(t-1)}$	0.31** (8.03)	0.30** (7.70)	0.28** (7.54)	0.29** (6.77)
$\text{dFTA}_{ij,(t-1)-(t-2)}$		0.22** (4.80)	0.19** (4.24)	0.16** (2.62)
$\text{dFTA}_{ij,(t-2)-(t-3)}$			0.14** (2.34)	0.09 (1.12)
$\text{dFTA}_{ij,(t+1)-t}$				−0.06 (−1.02)
Constant	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)
Total ATE[a]	0.31	0.52	0.61	0.45
Overall R^2	0.0007	0.0009	0.0008	0.0006
No. observations	36,563	34,105	31,172	24,642

t-statistics in parentheses. The dependent variable is the (natural log of the) real bilateral trade flow. * (**) denotes statistical significance at 5 (1) percent level in one-tailed t-test. Coefficient estimates for country-and-time effects are not reported for brevity.

[a] Total ATE is the sum of the statistically-significant FTA coefficient estimates.

request using trade flows scaled by the product of real GDPs). We start by first-differencing the natural logarithm of X_{ijt}, creating $\text{dln}X_{ij,t-(t-1)}$. Second, we regress $\text{dln}X_{ij,t-(t-1)}$ on 768 country-and-time effects ($\text{Dum}_{i,t-(t-1)}$, where i denotes a country and $t-(t-1)$ a 5-year period, e.g., 1995–2000) and retain the residuals. Third, we difference FTA_{ijt}, creating $\text{dFTA}_{ij,t-(t-1)}$, and regress $\text{dFTA}_{ij,t-(t-1)}$ on the same 768 country-and-time fixed effects and retain these residuals. Fourth, a regression of the residuals from the first (dln$X$) regression on the residuals from the second (dFTA) regression will yield unbiased estimates of the ATE effect of an FTA holding constant time-varying multilateral price terms.

The procedure described above is equivalent to estimating:

$$\begin{aligned}\text{dln}X_{ij,t-(t-1)} = {} & \beta_6 \text{dFTA}_{ij,t-(t-1)} + \beta_{i,t-(t-1)}\text{Dum}_{i,t-(t-1)} + \beta_{j,t-(t-1)}\text{Dum}_{j,t-(t-1)} \\ & + v_{ij,t-(t-1)} \end{aligned} \tag{12}$$

where $v_{ij,t-(t-1)} = \varepsilon_{ijt} - \varepsilon_{ij,t-1}$ is white noise. With nine years in the panel, we have 8 time periods $t-(t-1)$. Since there are 96 countries that can potentially trade, our procedure above effectively introduces 768 (=8×96) country-and-time fixed effects ($\text{Dum}_{i,t-(t-1)}$ and $\text{Dum}_{j,t-(t-1)}$) to account for the changes in the unobservable theoretical multilateral resistance terms, $\text{dln}P^{1-\sigma}_{i,t-(t-1)}$ and $\text{dln}P^{1-\sigma}_{j,t-(t-1)}$, to obtain an unbiased estimate of β_6. In the context of the theoretical model, the 768 estimates of $\beta_{i,t-(t-1)}$ and $\beta_{j,t-(t-1)}$ can be interpreted as changes in the countries' multilateral resistance terms.

Table 6 reports the coefficient estimates for the effects of concurrent, lagged and future changes in FTA (dFTA) on trade flow changes. Column (1) reports the ATE without any lagged effects; the ATE is 0.30. Including one lagged change in FTA increases the cumulative ATE to 0.52. With two lagged values of dFTA, the cumulative ATE of an FTA is 0.61. This estimate implies that an unbiased ATE of an FTA on the trade between a pair of countries is *84%*.[23] This

[23] Additional lags of FTA were also evaluated. None of these extra lags were economically or statistically significant. The results suggest that FTAs have a *level* effect on trade, but do not influence the long-run growth rate of trade. Bayoumi and Eichengreen (1997) used first-differenced data to estimate the effects on member trade of the six-member EEC and seven-member EFTA, but did not account for theoretically-motivated time-varying multilateral price resistance terms and only examined one 15-year period for these two agreements (1957–1972).Yet, the implied cumulative ATE for the EEC was 0.60, very similar to our estimated ATE of 0.61.

result is only slightly smaller than the comparable fixed-effects estimate in column (4) of Table 5 (0.76). For completeness, column (4) of Table 6 provides the results of adding a future change in FTA. Analogous to the fixed-effects results, future changes in FTA have no significant effect on trade flow changes. Thus, our estimates using differenced panel data suggest that — after 10 years — an FTA essentially *doubles* the level of members' international trade, confirming the previous section's results.

6. Conclusions

The purpose of this paper was to answer the question posed in the title: Do free trade agreements actually increase members' international trade? A motivation for this paper was that — after forty years of gravity equation estimates of the (treatment) effect of FTAs on trade flows — there seems no clear and convincing empirical evidence using the workhorse for empirical international trade studies that the answer is "yes." This seems surprising in light of the proliferation of FTAs in the last 15 years and widespread expectations that such agreements should increase trade. Our goal has been to provide a thorough empirical analysis of the average treatment effect of FTAs on trade, in light of prevailing knowledge on the theoretical foundations for the trade gravity equation and on modern econometric techniques to estimate "average treatment effects."

Our answer to the question: *yes!* Standard cross-section techniques using instrumental variables and control functions do not provide stable estimates of these ATEs in the presence of endogeneity and tests of overidentifying restrictions generally fail. However, we find convincing empirical evidence using panel data of unbiased estimates of ATEs ranging from 0.61 to 0.76. These estimates are five to six times those using OLS. The average of these two ATE estimates is 0.685. Stated succinctly, this estimate suggests that an FTA will on average increase two member countries' trade about *100%* after 10 years (using $e^{0.685} = 1.98$). This is seven times the effect estimated using OLS ($e^{0.13} = 1.14$, or 14% increase).

Some caveats are necessarily in order. While we have addressed multilateral price (resistance) terms of a country pair, we have not addressed the general equilibrium "comparative static" effects of an FTA on two members' trade. Nor have we addressed the impact of such agreements on trade with nonmembers, nor trade among nonmembers. Moreover, we have not addressed the welfare implications of FTAs. These are topics left for other research; our focus has been solely on trying to provide policymakers with more resolution on an unbiased estimate of the average treatment effect of an FTA on two members' trade. A second caveat is that the effect of an FTA is likely to differ depending upon the agreement. We consider this a useful extension for future research. A related caveat to the previous one is that the effect of an FTA on two members' trade may be influenced by the economic size, per capita incomes, and even distance between the two countries. That is, there may be interactive effects not addressed in this study, to limit its scope. This topic is left for future research.

Nevertheless, in light of the wide range of previous estimates of the average treatment effect of FTAs on trade, we hope that future research will acknowledge the importance of the endogeneity of FTAs, as the international trade literature recognized years ago for multilateral trade volumes and countries' tariff and nontariff barrier policies. Researchers using the gravity equation have already started to account for the endogeneity of multilateral price terms. We hope that future research will also come to appreciate the need to adjust also for endogeneity of FTAs when estimating the effects of such agreements on trade flows.

Acknowledgements

The authors are grateful to Jonathan Eaton, Judith Dean, Rob Feenstra, Scott Kastner, Mika Saito, Curtis Simon, Jeff Wooldridge, Eduardo Zambrano, two anonymous referees and participants at presentations at the American Economic Association annual meetings, Midwest International Economics meetings, U.S. International Trade Commission, and ISNIE meetings for excellent comments on earlier drafts. We thank Doru Cojuc, Katie Martin, Kushlani Rajapakse, Duncan Stewart, Li Zhang and Mike Jaszkowiak for excellent research assistance. The authors are very grateful for financial support to the National Science Foundation under grant numbers SES-0351018 (Baier) and SES-0351154 (Bergstrand), the Center for International Trade at Clemson University, and the Mendoza College of Business and the Kellogg Institute for International Studies at the University of Notre Dame.

Data appendix

The following is a list of the 96 countries potentially used in the regressions, depending upon availability of non-zero and non-missing trade flows:

Austria	Belgium–Luxembourg	Denmark
Finland	France	Germany
Greece	Ireland	Italy
Netherlands	Norway	Portugal
Spain	Sweden	Switzerland
United Kingdom	Canada	Costa Rica
Dominican Republic	El Salvador	Guatemala
Haiti	Honduras	Jamaica
Mexico	Nicaragua	Panama
Trinidad and Tobago	United States	Argentina
Bolivia	Brazil	Chile
Colombia	Ecuador	Guyana
Paraguay	Peru	Uruguay
Venezuela	Australia	New Zealand
Bulgaria	Hungary	Poland
Romania	Egypt	India
Japan	Philippines	Thailand
Turkey	Korea	Algeria
Angola	Ghana	Kenya
Morocco	Mozambique	Nigeria
Tunisia	Uganda	Zambia
Zimbabwe	China (Hong Kong)	Indonesia
Iran	Israel	Pakistan
Singapore	Sri Lanka	Syrian Arab Republic
China,P.R.: Mainland	Albania	Bangladesh
Burkina Faso	Cameroon	Cyprus
Côte d'Ivoire	Ethiopia	Gabon
Gambia, The	Guinea–Bissau	Madagascar
Malawi	Malaysia	Mali
Mauritania	Mauritius	Niger
Saudi Arabia	Senegal	Sierra Leone
Sudan	Congo, Dem. Rep. of	Congo, Republic of

References

Abrams, R.K., 1980. International trade flows under flexible exchange rates. Federal Reserve Bank of Kansas City. Economic Review 65 (3), 3–10.

Aitken, Norman D., 1973. The effect of the EEC and EFTA on European trade: a temporal cross-section analysis. American Economic Review 5, 881–892 (December).

Anderson, James E., 1979. A theoretical foundation for the gravity equation. American Economic Review 69 (1), 106–116 (March).

Anderson, James E., van Wincoop, Eric, 2003. Gravity with gravitas: a solution to the border puzzle. American Economic Review 93 (1), 170–192 (March).

Baier, Scott L., Bergstrand, Jeffrey H., 2001. The growth of world trade: tariffs, transport costs, and income similarity. Journal of International Economics 53 (1), 1–27 (February).

Baier, Scott L., Bergstrand, Jeffrey H. "On the endogeneity of international trade flows and free trade agreements." Manuscript, August 2002, http://www.nd.edu/~jbergstr/Working_Papers/EndogeneityAug2002.pdf.

Baier, Scott L., Bergstrand, Jeffrey H., 2004a. Economic determinants of free trade agreements. Journal of International Economics 64 (1), 29–63 (October).

Baier, S.L., Bergstrand, J.H. "Do free trade agreements actually increase members' international trade?" Manuscript, October 2004b, http://www.nd.edu/~jbergstr/Working_Papers/BaierBergstrandFTA2Oct2004.pdf.

Bayoumi, Tamim, Eichengreen, Barry, 1997. Is regionalism simply a diversion? Evidence from the evolution of the EC and EFTA. In: Ito, Takatoshi, Krueger, Anne O. (Eds.), Regionalism vs. Multilateral Arrangements. The University of Chicago Press, Chicago.

Bergstrand, Jeffrey H., 1985. The gravity equation in international trade: some microeconomic foundations and empirical evidence. Review of Economics and Statistics 67 (3), 474–481 (August).

Bergstrand, Jeffrey H., 1989. The generalized gravity equation, monopolistic competition, and the factor-proportions theory in international trade. Review of Economics and Statistics 71 (1), 143–153 (February).

Bergstrand, Jeffrey H., 1990. The Heckscher–Ohlin–Samuelson model, the Linder hypothesis, and the determinants of bilateral intra-industry trade. Economic Journal 100 (4), 1216–1229 (December).

Boisso, Dale, Ferrantino, Michael, 1997. Economic distance, cultural distance, and openness in international trade: empirical puzzles. Journal of Economic Integration 12 (4), 456–484 (December).

Brada, Josef C., Mendez, Jose A., 1985. Economic integration among developed, developing and centrally planned economies: a comparative analysis. Review of Economics and Statistics 67 (4), 549–556 (November).

Cheng, I.-Hui, Wall, Howard J., 2002. Controlling for heterogeneity in gravity models of trade. Federal Reserve Bank of St. Louis Working Paper, vol. 1999-010C.

Deardorff, Alan, 1998. Does gravity work in a neoclassical world? In: Frankel, Jeffrey A. (Ed.), The Regionalization of the World Economy. The University of Chicago Press, Chicago.

Eaton, Jonathan, Kortum, Samuel, 2002. Technology, geography, and trade. Econometrica 70 (5), 1741–1779 (September).

Egger, Peter, 2000. A note on the proper econometric specification of the gravity equation. Economics Letters 66, 25–31.

Eichengreen, Barry, Irwin, Douglas, 1995. Trade blocs, currency blocs and the reorientation of world trade in the 1930. Journal of International Economics 38 (1/2), 1–24 (February).

Feenstra, Robert, 2004. Advanced International Trade: Theory and Evidence. Princeton University Press, Princeton, NJ.

Felbermyer, Gabriel J., Kohler, Wilhelm. "Exploring the intensive and extensive margins in world trade." Manuscript, August 2004.

Frankel, J.A., 1997. Regional Trading Blocs. Institute for International Economics, Washington, DC.

Frankel, Jeffrey A., Romer, David, 1999. Does trade cause growth? American Economic Review 89 (3), 379–399 (June).

Frankel, Jeffrey A., Stein, Ernesto, Wei, Shang-Jin, 1995. Trading blocs and the Americas: the natural, the unnatural, and the super-natural. Journal of Development Economics 47 (1), 61–95 (June).

Ghosh, Sucharita, Yamarik, Steven, 2004. Are regional trading arrangements trade creating? An application of extreme bounds analysis. Journal of International Economics 63 (2), 369–395 (July).

Gilpin, Robert, 2000. The Challenge of Global Capitalism. Princeton University Press, Princeton, NJ.

Heckman, James J., 1997. Instrumental variables: A study of implicit behavioral assumptions used in making program evaluations. Journal of Human Resources 32 (3), 441–462 (Summer).

Heckman, James J., 2001. Micro data, heterogeneity, and the evaluation of public policy: Nobel lecture. Journal of Political Economy 109 (4), 673–748 (August).

Helpman, Elhanan, Krugman, Paul, 1985. Market Structure and Foreign Trade. MIT Press, Cambridge, MA.

Lawrence, Robert Z., 1996. Regionalism, Multilateralism, and Deeper Integration. The Brookings Institution, Washington, DC.

Lawrence, Robert Z., 1998. Comment on 'The role of history in bilateral trade flows'. In: Frankel, Jeffrey A. (Ed.), The Regionalization of the World Economy. The University of Chicago Press, Chicago, pp. 57–59.

Lee, Jong-Wha, Swagel, Phillip, 1997. Trade barriers and trade flows across countries and industries. Review of Economics and Statistics 79, 372–382.

Linnemann, Hans, 1966. An Econometric Study of International Trade Flows. North-Holland, Amsterdam.

Maddala, G.S., 1983. Limited-Dependent and Qualitative Variables in Econometrics. Cambridge University Press, Cambridge, MA.

Magee, Chris, 2003. Endogenous preferential trade agreements: an empirical analysis. Contributions to Economic Analysis and Policy, vol. 2, no. 1. Berkeley Electronic Press.

McLaren, John E., 1997. Size, sunk costs, and Judge Bowker's objection to free trade. American Economic Review 87 (3), 400–420 (June).

Oguledo, Victor I., MacPhee, Craig R., 1994. Gravity models: a reformulation and an application to discriminatory arrangements. Applied Economics 26, 107–120.

Preeg, Ernest H., 1998. From Here to Free Trade. The University of Chicago Press, Chicago.

Redding, Stephen, Venables, Anthony J., 2000. Economic geography and international inequality. Centre for Economic Policy Research Discussion Paper, vol. 2568.

Rose, Andrew, 2004. Do we really know that the WTO increases trade? American Economic Review 94 (1), 98–114 (March).

Rose, Andrew, van Wincoop, Eric, 2001. National money as a barrier to international trade: the real case for currency union. American Economics Review 91 (2), 386–390 (May).

Sapir, Andre, 1981. Trade benefits under the eec generalized system of preferences. European Economic Review 15, 339–355.

Tinbergen, Jan, 1962. Shaping the World Economy. The Twentieth Century Fund, New York.

Tomz, Michael, Judith Goldstein, Douglas Rivers, June 2004. "Membership has its privileges: the impact of GATT on international trade," Stanford University, Dept. of Political Science Working Paper.

Trefler, Daniel, 1993. Trade liberalization and the theory of endogenous protection: an econometric study of U.S. import policy. Journal of Political Economy 101 (1), 138–160 (February).

Wooldridge, Jeffrey M., 2000. Introductory Econometrics. South-Western College Publishing, Cincinnati, OH.

Wooldridge, Jeffrey M., 2002. Econometric Analysis of Cross-Section and Panel Data. MIT Press, Cambridge, MA.

World Bank. *World development indicators*, CD-ROMs. Washington, DC: World Bank, 2003.